Representative Compensation and Disability Claimant Outcomes*

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Abstract

Many claimants of Social Security Disability Insurance (SSDI) retain legal representation to help with the approval process. The Social Security Administration imposes strict rules on representative compensation. Representatives are only paid if claimants are awarded disability, and they are paid the lesser of 25 percent of the claimant's past due benefits or a pre-specified maximum fee ($6,000 since 2009). Because past due benefits are a function of the number of months claimants wait to be awarded, representatives face incentives to delay case resolution until past due benefits push the representative fees past the fee ceiling. We use difference-in-differences to evaluate how these incentives impact SSDI claimant wait times. After the fee ceiling increased in 2002, average wait times increased by 0.85 months among claimants for whom the fee threshold is more binding, implying a 2.6-5.6 month increase for claimants with representatives. This indicates that the structure of representative compensation does matter for case outcomes, and highlights the importance of interactions with auxiliary agents so common in modern social programs.

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1 Introduction

A vast literature explores how the structure of various programs affects participants' incentives and behavior. However, program structure can also affect the incentives and behavior of administrators and intermediaries who are not direct beneficiaries of the program. There is far less work documenting how the incentives of these gatekeepers to social programs, such as caseworkers, judges, or legal representatives, affect beneficiaries' outcomes. These indirect effects can influence the overall effectiveness and efficiency of the program. As a case study, we document such patterns in one particular setting: legal representatives in the Social Security Disability Insurance (SSDI) program.

SSDI is one of the largest components of the US safety net in terms of dollars, transferring roughly $11.7 billion a month to approximately 8.4 million disabled workers and their families (Social Security Administration, 2020). Since 2006, between 2 and 3 million disabled workers have applied for SSDI each year (Social Security Administration, 2020). The SSDI application process is complicated and drawn-out; in July 2019, 13.5 percent of 2018 claimants were still waiting a final decision and there were over 5,000 claimants from 2012 waiting for a final decision (Social Security Administration, 2020). This process leads many claimants to obtain legal representation. Each year, nearly 400,000 representative payments are made by the Social Security Administration (SSA) for SSDI claimant cases, resulting in approximately $1.1 billion of benefits being paid directly to representatives. The prevalence of representation has also become more common over time (see Figure 1). Between 2001 and 2018 the number of representative payments relative to awards increased from 27 to 55 percent, suggesting that over one half of successful cases have representation.[1]

The SSA imposes a unique representation fee schedule, which restricts representative compensation to the lesser of 25 percent of the claimant's past due benefits or a prescribed fee ceiling, which has been $6,000 since 2009, but just increased to $7,200 in November

[1]Hoynes et al. (2021) suggest that in 2014, nearly 15 percent of all initial SSDI claims and between 2007 and 2014 82 percent of appealed claims (regardless of case outcome) obtained representation.

2022. Because past due benefits are increasing in the number of months claimants wait to be awarded, this schedule decreases representatives' incentives for timely case resolution. By design, representative pay increases as wait times increase until 25 percent of expected past due benefits crosses the $6,000 threshold. After this point, the marginal revenue to the representative of further delays is zero. These incentives may impact the amount of time it takes for applications and appeals to be resolved, and by consequence, a host of claimant outcomes (Autor et al., 2017; Coe et al., 2014; Michaud et al., 2019).

Even though over $1 billion of disability claimant benefits are transferred to representatives each year, we have no evidence on how the representative fee structure affects claimant outcomes.[2] We are able to provide the first evidence of how representative fee caps affect claimant wait times. In this paper we use a difference-in-differences design to determine if the representative fee structure leads to longer wait times. To do this we use an underutilized dataset, the administrative Disability Analysis File (DAF) Public Use File (PUF), a 10 percent random sample of all SSDI recipients' benefit histories between 1995 and 2018. This allows us to explore detailed variation in claimant payments and wait times, in ways that are not possible with smaller surveys. Although the DAF PUF does not include information on representation, we exploit changes in the representative fee schedule to identify the reduced form effect on claimant wait times. Using difference-in-differences methods, we test whether a large increase in the representative fee cap in 2002 lead to longer wait times, particularly for individuals with earnings-levels that make the fee cap relevant relative to workers with earnings that make the cap less binding. We find that the increase in the cap in 2002 increases wait times for more-affected workers by about 0.85 months, on average. Using the representative payment-to-award ratio between 2002 and 2009 of 32 percent to proxy for the representation rate would imply, under some assumptions, that wait time increased by 2.6 months for those with representation. This strategy sheds light on how we should expect the 2022 fee cap increase to affect wait times.

[2]A concurrent working paper (Hoynes et al., 2021) documents how retaining representation affects case outcomes, but it does not explore how features of the fee schedule affect outcomes.

These estimates are robust to controlling for diagnostic code fixed effects and restricting the sample to claimants with more similar benefit levels. One concern is that these shifts in claimant wait time are due to the economic recession in 2001, not the 2002 policy change. However, the estimates are robust to controlling for gender by education by region specific labor market conditions and we do not find a relationship between how hard a region was hit by the 2001 recession and our estimated effect of the fee increase in that region. Moreover, we find a sharp, discontinuous increase in 2002 and lasting impacts on wait times, which are more consistent with the permanent fee increase rather than transitory economic shocks or slow moving trends in backlogs that accompanied the 2001 recession. Despite similar concerns about local economic conditions in 2009, we find a similar, albeit smaller, effect on wait times associated with a smaller fee cap increase that occurs in 2009.

To the extent possible, we also document potential mechanisms. Although the fee ceiling increase could affect both representative effort and attention to timeliness, we do not see systematic shifts in the award rate, number of awards, awardee characteristics, or the propensity of early onset dates, which we might expect if representatives put in more effort, leading to higher award rates. As increased wait times are driven by actual delays and not an increased propensity to back-date onset dates, the pattern is more consistent with a reduced focus on timeliness rather than an increase in effort. We see distributional changes in claimant wait time that are consistent with strategic behavior. Namely, we document that a fee-maximizing representative will increase wait times the most for claimants in the middle of the monthly entitlement distribution, and we find that those claimants do have the largest increase in wait times as a result of the fee increase.[3]

Taken together, this evidence suggests that the incentives created by the SSDI representative fee structure do lead representatives to slow down case processing, leading to higher fee payments for the representatives and longer wait times for the claimants. Many disability

[3] We do not observe sharp bunching in the wait time distribution, likely due to representatives not having perfect control over every step in the application process. Appendix B explores these distributional changes in wait time in more detail.

claimants are liquidity constrained (Deshpande et al., 2021), suggesting longer wait times could harm claimants. Wait times from disability onset to benefit allowance are a critical piece of the disability claim process and are of recent interest to policymakers. As recently as 2019, several US senators introduced a bill specifically aimed at reducing SSDI wait times, citing the statistic that in 2017, more than 10,000 people died while waiting for their SSDI benefits to begin.[4] Autor et al. (2017) find that longer wait times lead to a reduction in post-decision labor supply, and Coe et al. (2014) show that while claimants wait for their SSDI decision they increase their reliance on SNAP benefits.[5] Other work on SSDI highlights that increasing the ordeal costs of applying decreases the targeting efficiency of the program, another potential consequence of long wait times (Deshpande and Li, 2019; Foote et al., 2018).

Although there is significant work exploring the impacts of disability insurance programs like SSDI on claimant outcomes, such as labor supply (French and Song, 2014; Gelber et al., 2017; Maestas et al., 2013), financial stability (Deshpande et al., 2021), and household responses (Autor et al., 2019), there is very little work examining how the structure of SSDI affects other agents that participate in the process. We are only aware of one other paper that examines the role of SSDI representatives (Hoynes et al., 2021). Hoynes et al. (2021) bring unique administrative SSA data to bear on the distinct and important question of whether having a representative improves case outcomes for claimants relative to not having a representative at all. Using a novel identification strategy that leverages geographic variation in disability representative markets, they find that retaining representation (versus not retaining representation) increases the probability of allowance in the initial review. We explore the distinct question of how representatives respond to the SSA imposed fee structure. It is possible for representatives to both improve the probability of being awarded disability relative to a claimant with no representation and adjust at the margin in response

[4]S. 2496 (116th): Stop the Wait Act.

[5]Michaud et al. (2019) document a negative relationship between county-level wait times in one year and the number of applications in the next, suggesting that long wait times might discourage future applications.

to payment incentives.

We provide the first evidence on how the SSA representative fee schedule affects claimant outcomes. Prior to Hoynes et al. (2021), there was very little information on who obtains representation and how they fare in the disability process. Even their data only begins in 2010, after the last fee ceiling increase, so they cannot evaluate how the structure of the representative fee schedule affects case outcomes.[6] As we discuss below, data limitations in the DAF PUF restrict us to reduced form estimates and preclude us from examining client selection among representatives, but nevertheless we are able to provide the first, crucial evidence on how SSA representative gatekeepers respond to incentives from the fee schedule.

In addition to the specific insights we provide about legal representation in the SSDI system, we add to a growing literature documenting how incentives faced by programmatic gatekeepers like doctors, judges, attorneys, or caseworkers affect treatment and outcomes of individuals (Cabral and Dillender, 2022; Clemens and Gottlieb, 2014; Fitzpatrick et al., forthcoming; Yang, 2016). We also contribute to the literature that studies the connection between attorney compensation and client outcomes. We study a misalignment in attorney and client incentives in a civil setting to test whether representatives act in their clients' best interests. Recent work on public defenders has shown that the structure of the compensation set by the public defender's office (flat fee vs hourly) significantly impacts pleas and convictions in criminal cases (Agan et al., 2021; Lee, 2021). Our paper adds new evidence by estimating the response of legal representation to compensation incentives in an important area of civil law that interacts with a crucial safety net program in the US.

Finally, this paper connects to an active policy debate about reducing the administrative hurdles and delays inherent to accessing social programs in the US. Researchers and policy analysts have examined these burdens in other domains, including Medicaid, unemployment benefits, and SNAP (Herd and Moynihan, 2020). Recent executive orders aim to curtail such hurdles for certain social security payments and disaster relief.[7] Ours is the first paper to

[6]We further discuss similarities and differences with Hoynes et al. (2021) in section 7.

[7]See: `https://www.bloomberg.com/news/articles/2021-12-13/biden-aims-to-slash-time-tax-`

document this phenomenon in the context of legal representation for disability application. We highlight how the structure of SSDI representative compensation can affect claimant wait times, a critical outcome on its own and a determinant of further downstream outcomes like future labor supply (Autor et al., 2017).

2 The SSDI Application Process

SSDI is a social insurance program, designed to insure against the risk that an individual becomes disabled and is no longer able to work and financially support themselves. To be eligible, workers must earn sufficient Social Security work credits in covered jobs and have sufficient work in the past 10 years. Workers must also document that they have a permanent disability that will (1) last at least one year, (2) keep them from performing the work they did previously, and (3) prohibit them from adjusting to other work. Disabled workers who apply to SSDI will first have their application reviewed at a local SSA field office, to verify the individual meets all non-medical requirements. If they do, the application is passed on to a state-level Disability Determinations Services (DDS) agency. A DDS examiner will review the individual's case, evaluate available evidence, or seek additional evidence through a consultative examination. If it is determined that the worker is disabled, the case is returned to the field office where benefits are calculated and payments begin. Between 33 and 35 percent of initial applicants are allowed at this point (Social Security Administration, 2020). If the examiner determines the worker does not meet the disability requirements, the application is sent back to the field office, allowing the claimant to appeal the decision.

A claimant has 60 days to appeal a negative decision, in which case it is sent back to DDS to a new examiner for Reconsideration. Only 2-3 percent of initial applicants are allowed through Reconsideration (Social Security Administration, 2020). If the applicant is again denied, they can appeal the decision again and it will then be sent to the appropriate hearing office, where it will be heard by an administrative law judge (ALJ). The ALJ will hold a

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hearing with the individual and any legal representation the individual has retained. The claimant can present further evidence regarding the severity of the disability, the onset date of the disability, or the limitations caused by the disability. Between 15 and 19 percent of initial applicants are allowed through an ALJ decision, meaning 31-36 percent of all recipients are approved by an ALJ (Social Security Administration, 2020).[8] If a claimant is denied by an ALJ, they can appeal again to the SSA Appeals Council then for Federal Review. Very few cases are pursued to these final stages.

3 Disability Representation Compensation and Incentives

At any point in the process a claimant can retain legal representation. Both attorneys and non-attorneys[9] can act as representatives. Estimates suggest that in recent years, around 83 percent of all cases with a representative have an attorney representative, but non-attorneys were not eligible to receive compensation until 2010, after the fee cap increases (Hoynes et al., 2021). There is only limited data on the utilization of representation in SSDI appeals. Using SSA administrative records, Hoynes et al. (2021) find that about 8 percent of initial claims retained legal representation in 2010. By 2014, this had increased to 15 percent. Aggregated representation payout data suggests 377,054 representative payments were made in 2018, relative to 761,481 total awards (48 percent) (Social Security Administration, 2019). On average over $1.1 billion a year of claimant benefits are paid directly to representatives for their services (Hoynes et al., 2021).

Just four years after the 1935 Social Security Act was passed, it was amended, giving SSA the ability to prescribe the maximum fee that can be charged for representation services (Section 206 of Social Security Act Amendments of 1939). This section was added to protect claimants from potential, predatory behavior and also gave SSA the ability to suspend

[8]Because the hearing process take a long time, these are the estimates between 2005-2013, to avoid censored data.

[9]To be eligible, non-attorneys must hold a bachelor's degree, pass a written exam administered by SSA, obtain liability insurance, undergo a criminal background check, and complete continuing education courses (see Social Security Act, Section 206 [42 U.S.C.406].

persons from providing representation if they did not adhere to the SSA representative policies. In 1990, Section 206 was further adjusted specifying the 25% or maximum fee rule (originally $4,000) to speed up the attorney fee payment approval process for most cases (Federal Register, 1990). It also clarified that SSA must approve all representation fee agreements. As such, claimants and representatives cannot legally agree to a fee outside the purview of SSA.[10]

SSA has outlined two important conditions for SSDI representative compensation. SSDI representatives (1) will only be compensated if the claimant is awarded disability insurance, and (2) will be paid a scheduled fee, which both the representative and the claimant must agree to, that does not exceed the lesser of 25 percent of past due benefits, or a prescribed dollar amount (Federal Register, 2009). This maximum amount has changed over time. The Omnibus Budget Reconciliation Act of 1990 specified the SSDI representative compensation formula, with a maximum prescribed amount of $4,000. This limit was increased to $5,300 on January 17, 2002, and then again to $6,000 on June 22, 2009 (announced January 29, 2009). These fee increases are authorized by the SSA Commissioner, as long as they do not exceed the cumulative PIA cost of living adjustments since 1991. The new fee limits are published in the Federal Register Notice. There is not much recorded about the lead up to these fee increases. However, the National Organization of Social Security Claimant Representatives indicates that prior to the 2002, 2009, and 2022 changes they spent several year pushing for a change, suggesting there was significant lobbying from representatives.[11] Fee agreements that are approved by the SSA after these dates are eligible for the higher fee threshold.[12]

[10]Under certain circumstances, representatives can file a fee petition, asking for a higher payment. This is mostly restricted to settings where there was no initial fee agreement, or SSA did not approve the initial fee agreement, but the petition must still be approved by SSA.

[11]See `https://nosscr.org/acting-commissioner-credits-nosscr-for-first-fee-cap-raise-in-13-years/`.

[12]Since a claimant can seek representation at any point in the process, this date does not necessarily correspond to other dates in the SSDI process, such as the disability onset or application date. Under certain circumstances, representatives can file a fee petition, asking for a higher payment. This is mostly restricted to settings where there was no initial fee agreement, or SSA did not approve the initial fee agreement.

Crucial to this compensation schedule are past due benefits. Upon SSDI allowance, recipients are eligible for past due benefits from the month they became "entitled" for SSDI to the month they were awarded benefits. Because of the required five month waiting period, the entitlement date is typically five months after the established disability onset date.[13] Throughout, we will refer to the first month the individual is entitled to benefits as the month of entitlement. Past due benefits are the claimant's monthly benefit level multiplied by the number of full months between the entitlement date and award date. Monthly benefits are calculated using the same SSA Averaged Indexed Monthly Earnings (AIME) and Primary Insurance Amount (PIA) formulas used to calculate social security old-age and retirement benefits. The AIME is your average monthly earnings in your 35 years of highest earnings, after indexing for inflation. This number is then plugged into the progressive PIA formula with a 90 percent replacement rate for an AIME under \$996 (2021 values), a 32 percent replacement rate for an AIME between \$996 and \$6,002, and a 15 percent replacement rate for an AIME over \$6,002. Thus, for pre-determined earnings history and benefit level, there is a number of months of "wait time" from entitlement to award date that will push 25% of past due benefits over the fee ceiling. At this point, the representative's fee is maxed out and no longer dependent on wait time or past due benefits.

To understand how this compensation structure affects the representative's incentives, it is useful to consider the representative's objective function. We assume the representative is profit maximizing. As such, the conditional compensation stipulation incentivizes representatives to take on claimants that are likely to be awarded disability insurance and to push for an award. Given data constraints we are not able to examine this decision and will focus on the second feature of the payment structure.[14]

The fee structure stipulation creates incentives for disability representatives to allow cases

[13] In partially favorable awards, where the judge establishes a later onset date, the entitlement date might not be five months after the disability onset date, but still corresponds to the first month the claimant is entitled to benefits.

[14] We leverage public-use Census data to examine changes in the probability of receiving SSDI across the income distribution. We do not find any evidence that people for whom the fee cap is more binding are thus more likely to receive SSDI after the fee increase in 2002.

to move slowly through the decision process–which increases past due benefits and representative compensation–until 25 percent of past due benefits exceeds \$6,000 (or the prevailing maximum threshold). Conditional on taking the case, the representative's expected revenue depends on both the probability of winning the case and the expected wait time as follows

$$E(revenue_{it}(e,\tau,X_i)) = \rho(e,X_i)\min\{0.25 * b(X_i) * E(Wait\ Time(e,\tau,X_i)), FeeCap_t\} \quad (1)$$

Expected revenue is a function of the representative effort (e), focus on timeliness (τ), and the claimant's characteristics (X_i). The probability of winning, $\rho(e, X_i)$, depends on representative effort and claimant characteristics; $b(X_i)$ is individual i's benefit level (which depends on their work history and other characteristics). Wait time is a function of claimant characteristics, but representative actions could also affect wait time (and expected revenue if wait time is below the prevailing fee ceiling). Importantly, claimants and disability representatives do not have perfect control of the application timeline. For example, during the DDS review or ALJ appeal much of the wait time is driven by SSA action, not claimant or representative behavior. However, after each stage of SSA evaluation (i.e., the initial review, Reconsideration, hearing, appeals council, and final federal evaluation) claimants are given 60 days (2 months) to appeal negative SSA decisions before the case is automatically closed. Claimants can use this time to gather more supporting evidence and decide if they want to appeal. This suggests there are several ways representatives could increase wait times.

First, they can put more effort (e) into the case by collecting more evidence and building a more convincing argument, which might take time. Representatives might be willing to increase effort in this way if it increases ρ, the probability of winning the case. Second, they can put less focus on timeliness (τ). For example, rather than working quickly to resubmit missing information or submit an appeal, they can de-prioritize the case, submitting it closer to the 60 day deadline. There are four steps in the appeal process when this could happen (although most cases will only undergo two of them), suggesting 4 to 8 months of potential delay time. This change in prioritization and timeliness need not be nefarious, but since there is not pressure to move the case along quickly and in fact incentives to delay, representatives might move their focus to other demands. This reduction in timeliness is unlikely to affect the probability of winning, except in extreme circumstances where a deadline is missed and the case is closed. It can however increase wait times, past due benefits, and the representative's expected pay off. The reduction in timeliness is also likely to reduce psychic costs for the representative, as they relax their time constraints.[15]

Although both channels are potentially important, the monetary incentives associated with reducing timeliness will play a larger role in responses to the fee cap increase. If effort increases the probability of winning, the incentive to increase effort always exists. Proximity to the fee cap only marginally affects this since the payout for winning is large and positive, both before and after the increase in the fee cap. However, the return to reducing timeliness jumps dramatically for cases near the pre-existing cap once the cap increases. We argue that representative responses to the change are more likely a result of this large increase in the returns to reducing timeliness rather than the small increase in returns to increasing the probability of winning.

For example, consider a simplified case where the individual's benefit level and expected wait time pushes past due benefits right to the pre-2002 cap of $4,000. Suppose that the baseline probability of winning the case is 0.6, meaning the representatives expected revenue is $0.6 * 4,000 = 2,400$. Now, suppose the representative can exert additional effort and increase the probability of winning by 10 percentage points. This additional effort might include things like compiling a more complete documentation of the individual's disability and might take additional time. As such, suppose the additional effort also increases expected past due benefits to $5,300. When exerting additional effort, the expected revenue is $0.7 * 4,000 = 2,800$ or $400 dollars higher, since the $4,000 cap is still in effect.

[15]This would suggest that reducing a focus on timeliness is both weakly revenue-increasing and cost-reducing.

Now, consider what happens when the fee ceiling is lifted to the 2002 level of $5,300. Under this scenario, the representatives expected revenue when exerting additional effort increases to $0.7 * 5,300 = 3,710$, or $1,310 higher than expected revenue without any increase in effort. Note, however, that the representative could alternatively reduce their timeliness rather than increase effort. If the representative relaxes their concern for timeliness, increasing wait time and expected past due benefit to $5,300, the expected revenue is $0.6 * 5,300 = 3,180$. The return to increasing ρ by 10 percentage points is $530 under the new cap ($3,710-$3,180), which is only $130 higher than the return to increasing ρ under the old cap. For cases that were at the pre-existing cap, the return to decreasing timeliness under the new cap is $780. As such, the return to decreasing timeliness in response to the cap increase is approximately six to seven times larger than the return to increasing ρ. In order for the fee cap to change effort, the effort cost to the representative must be more than $400 but less than $530. Although it seems likely the effects are driven by a de-emphasis of timeliness, we will look for supporting evidence of each of these channels directly in the data.

These representative incentives have the potential to impact claimant welfare. Longer wait times could result in claimants spending longer periods without income, potentially leading to even worse health. Even though benefits accrued during this waiting period would be eventually paid out in a lump sum as past due benefits, this might reduce welfare if claimants face binding borrowing constraints (Deshpande et al., 2021). Longer wait times that result in more past due benefits and more dollars of representative compensation would also result in claimants paying a larger share of lifetime benefits to representatives. These incentives also affect claimants in the income distribution differently. Past due benefits are increasing in Average Indexed Monthly Earnings (AIME), meaning the incentive to delay case resolution would be stronger for low-income claimants, when it takes more months to reach the fee ceiling. Representatives can also increase past due benefits by arguing for an earlier onset date. Since, this would increase overall claimant benefits without increasing the wait time from application to award, we explore this channel directly. Anecdotally, there is some evidence of representative compensation rules influencing claimant outcomes. In a 2012 report (Social Security Advisory Board, 2012), field office and DDS employees allege that many professional representatives deliberately delay documentation and evidence to increase past due benefits and representative compensation.

4 Data

We use the 2015 and 2018 Disability Analysis File (DAF) Public Use File (PUF) to examine how the SSA representative fee schedule affects claimant outcomes. The DAF is a compilation of SSA administrative data from various sources including the Mastery Beneficiary Record of beneficiary enrollment and the 831 files collected when the disability determination was made. It includes longitudinal records for adult SSI and SSDI recipients who have received benefits at any point since 1996, including spouses and dependents that receive benefits.[16] The PUF is a 10 percent random sample of beneficiaries in the DAF and has a more limited set of variables to avoid disclosure risk.[17] For example, there is limited geographic information and dollar amounts are rounded.

There are two components to the DAF PUF. The first is a demographic file, which contains the cross-section of claimant attributes including things like date of birth, sex, primary and secondary disability diagnostic codes, and the claimants' application and entitlement date. The second component is a set of annual files, which include monthly records including claimant eligibility status, benefits due, and benefits paid. By combining payment and eligibility with the demographic file, we are able to construct the number of months between application and award and entitlement and award for each individual. We will call the time from entitlement to approval as "wait time". Importantly, if a claimant or representative

[16] Information for 10-17 year old recipients was added later, starting with beneficiaries that have received benefits after 2005.

[17] We combine the random samples from both the 2015 and 2018 files. As these are independent random samples, and since the population of claimants changes from year to year, there is a less than 1 percent change an individual might show up twice in the analysis sample.

can successfully argue that the disability onset date is earlier (and thus the entitlement date is earlier) this will increase our measure of wait time, so we think of this measure as capturing time from eligibility to award.[18] Using the first recorded monthly benefits owed to the individual, we can also measure the individual's initial benefit level. These measures allows us to construct the claimants expected past due benefits by multiplying the initial benefit level by the number of months between entitlement and award.

There are limitations in the DAF PUF. Most relevant to the question at hand, there is no record of whether legal representation was retained. As such, we can only identify the reduced form effects of the fee schedule on claimant outcomes. Unfortunately, this precludes us from doing a "placebo" test that shows there is no wait time effect among individuals without representation. However, we are able to provide "placebo" estimates by examining differential effects by diagnostic group, where some groups (like claimants with infections) are less likely to experience a binding fee cap and be affected by the policy change. To maintain the anonymity and privacy of SSDI recipients, dollar amounts are also rounded. Dollar values between 1 and 7 were rounded to 4, values between 8 and 999 were rounded to the nearest 10, and values between 1,000 and 49,000 are rounded to the nearest 100. SSA reports that average monthly benefits for workers is $1,257.65 (in 2019), suggesting our measured benefit level could be up to 50 dollars off. This will lead to measurement error in our construction of past due benefits and representative fees. Dollar values are also top coded for the top half percent of positive values. As past due benefits are large relative to monthly payments, past due benefit payments that are made to claimants are likely to be muted by top coding. However, we can calculate the expected past due benefits because we can infer the entitlement month, award month, and initial monthly benefit. Dates are also reassigned to the 15th of the month for privacy. As such, we observe the number of months of past due benefits owed with error. However, as only full months are counted towards past due benefits, our measure should be within one month of the actual number of months

[18] As we show in Appendix Figure A7, the changes in wait time in 2002 are not driven by an increase in the propensity of getting an earlier onset date.

benefits are owed. We also do not observe if an individual appeals a decision (and is thus more likely to retain representation).

Unfortunately, better data are not available to explore these relationships. Even the detailed administrative data Hoynes et al. (2021) use, only starts linking applicants to representative information starting in 2010, after the last maximum representative fee increase. Even with these data limitations, we provide the first estimates of how changes in the representative fee structure affects wait times for awardees.

We restrict the sample to only include SSDI recipients who applied in 1996 or later, so that we can accurately identify wait times. We also limit the sample to claimants whose application date was at least five years before the data collection date (either 2015 or 2018) to focus on resolved cases that are not truncated. We also focus on workers and exclude spouses and dependents (who cannot be linked to workers).[19] We make several other restrictions to address anomalies in the data and mitigate the influence of outliers. We exclude people whose application to entitlement period is over 25 months, people whose application to allowance period is over 60 months, and whose initial benefit level was over $2,300. Combined, these restrictions exclude only three percent of claimants in the DAF PUF.[20,21]

5 Empirical Strategy

We explore how the SSA representative fee schedule affects the appeal process and payments for claimants that are awarded disability. For each SSDI claimant, there is a kink in the

[19]For fee agreement calculations, past-due benefits include the monthly benefits credited for all auxiliary beneficiaries unless they have appointed their own representative. We cannot connect dependents to workers and we do not know if dependents appointed their own representative. As such, our measures of past-due benefits may be under counted, but proportional, as both worker and auxiliary benefits are a function of the worker's earning history. In the combined DAF-PUF samples from 2015 and 2018, 92.3 percent of recipients are workers, meaning that this is a concern in less than 8 percent of cases (some workers are likely to have multiple dependents).

[20]When including these outliers, we also find the fee change leads to an increase in wait times. The effect is qualitatively similar and statistically significant.

[21]Data are also collected for SSI recipients. However, the DAF PUF does not include sufficient information on SSI application, onset, and entitlement dates to construct accurate measures of wait time. We cannot examine how representative compensation affects outcomes of SSI only claimants.

potential representative fee, when the amount of time between the month of entitlement and allowance date pushes 25 percent of past due benefits over the current fee ceiling. Prior to this kink, the representative's marginal revenue for one month of waiting is 25 percent of the individual's monthly benefit. Once the threshold is crossed, the marginal revenue for an additional month of wait time is zero. Because past due benefits are a function of both claimant AIME and how many months lapse between the month of entitlement and allowance dates (the month before effectuation), the amount of time it takes for 25 percent of past due benefits to exceed the threshold will vary for each individual.[22] We will first explore how changes in the fee ceiling affect wait times for claimants that are more likely and less likely to face a binding fee threshold.

Past due benefits is the number of full months between the entitlement date and award date, multiplied by the claimant's monthly benefit. As such, we can back out the number of months from entitlement to award it would take for 25 percent of past due benefits to meet the current fee ceiling by solving the following equation

$$Monthly\ Benefit_i * Months * 0.25 = Threshold \tag{2}$$

This yields the number of months to reach the threshold, $Months = \frac{4 * Threshold}{Monthly\ Benefit_i}$. Because the monthly benefit varies across individuals this is an individual specific threshold, but the number of months to the threshold will fall as monthly benefits increase. For example, some individuals will reach the fee ceiling in less than a year, while others with lower benefit amounts might not reach the threshold until after 3 years of waiting. Backlogs in the disability determination system result in different baseline propensities that the individual will wait long enough for the maximum representative fee will be reached. As seen in Figure

[22] Past due benefits and past due benefits payable to the claimant are different. For SSDI claimants, past due benefits are used to calculate representative fees, past due benefits payable are the benefits claimants receive after subtracting direct payments to representatives, reductions due to SSI receipt, and prior overpayments (`https://secure.ssa.gov/poms.nsf/lnx/0203920032`). Past due benefits are calculated differently for SSI claimants. Interim Assistance Reimbursement for state pre-payments is deducted before representative fees are calculated. Also, past due benefits include the month of effectuation. In this paper we focus on SSDI claimants and SSDI past due benefits.

2, the threshold has increased twice, once in February 2002 by $1,300 (32.5 percent) and once in June 2009 by $700 (13.2 percent). An increase in the fee ceiling will lead to an increase in the number of months of past due benefits it takes to reach the fee threshold. In other words, representatives are compensated for more months of waiting. Because the fee ceiling is a dollar amount, it will become more binding as monthly benefit levels increase. As a result, the probability of reaching the fee ceiling increases almost linearly with benefit level.

Although the relationship between benefit level and whether the fee ceiling binds is monotonic, the relationship between benefit level and whether the claimant is affected by the fee ceiling *increase* is not necessarily monotonic. As shown in Figure 2, claimants with back pay from $0 to $15,999 will not be affected by the cap increasing in 2002.[23] Those claimants do not reach the cap when it is $4,000 or when it is $5,300. Similarly, although the cap increase in 2002 would increase the fee level for claimants with back pay greater than $21,199, the marginal fee for each additional dollar of back pay will remain zero, since these claimants will reach the cap under the $4,000 rule and under the $5,300 rule. There are far fewer claimants at these high benefit levels. The claimants who are affected are those for whom the cap binds prior to 2002 but, absent any changes in wait time, the new cap would not bind after 2002. Similar logic applies to the smaller cap increase in 2009. In theory, this could induce a quadratic relationship between benefit level and whether the claimant is affected if claimants with high benefit levels have long enough wait times, on average.

We explore this empirically in the first panel of Figure 3 by plotting the share of claimants in each $100 monthly benefit level bin who applied under the $4,000 cap and reached the $4,000 cap but who, based on their wait time, would not have reached the $5,300 cap. This provides a summary measure of the fraction of claimants at each benefit level who would be affected by the cap increase. Approximately 2.2 percent of claimants with benefit levels

[23]We calculate that approximately 54% of all awarded claimants have expected past due benefits greater than $0 but less than $15,999 and approximately 8% of all awarded claimants have expected past due benefits greater than $16,000 but less than $21,199.

from \$0-\$500 would be affected by the increase in 2002 whereas about 7.1 percent of those with benefit levels above \$500 would be affected. This figure exhibits a weak quadratic relationship, as fewer claimants with benefit levels above \$1,700 would be affected by the increase than those with benefit levels from \$500-\$1,700.

Panel 1 of Figures 3 motivates our difference-in-differences strategy. We combine this variation in the baseline propensity of reaching the fee ceiling driven by differences in earnings with the 2002 policy change in the fee ceiling. In particular, the relationship suggests that claimants with initial benefit levels between \$0-\$500 are less affected by the change in the fee ceiling than those with initial benefit levels above \$500. The intuition is that claimants with low benefit levels are unlikely to reach the fee ceiling in either period. As such, representatives have little incentive to change their behavior with regard to these cases after the threshold increases. Claimants with higher benefit levels, however, are likely to reach the fee ceiling prior to the increase but less likely to reach the fee ceiling once it's raised. As a result, representatives face a high return from lengthening those cases.

To formalize this, we categorize claimants with benefit levels from \$0-\$500 as the "less-affected" or "control" group and claimants with benefit levels above \$500 as the "more-affected" or "treated" group.[24] We then estimate two difference-in-differences models comparing less-affected to more-affected claimants before versus after the fee ceiling increases in 2002 and 2009. Both models use the following specification:

$$Outcome_{it} = \alpha_0 + \beta(More\text{-}Affected_i * PostIncrease_t) + \gamma More\text{-}Affected_i + \phi_t + X_i\Gamma + \varepsilon_{it} \quad (3)$$

When estimating the effect of the 2002 increase, we limit the sample to claimants who applied between 1996 and the first half of 2009 (plausibly under either the \$4,000 or \$5,300 cap). Then, we define $PostIncrease_t$ as equal to zero if the the claimant applied from January 1996 to February 2002 under the \$4,000 cap and equal to one if they applied from March 2002

[24]Over the full time period of our sample, 15% of claimants are in the "less-affected" group.

to June 2009 under the \$5,300 cap. $Outcome_{it}$ is either the claimant's wait time (in months) from entitlement to approval or the hypothetical representative's fee calculated using the claimant's past due benefits and the fee ceiling in effect in year t. $More\text{-}Affected_i$ is equal to zero if the claimant's monthly benefit level is \$0-\$500 and equal to one if their benefit level is above \$500. Finally, X_i is a matrix of individual-specific characteristics such as: sex, age, primary condition, and \$100 benefit level bin fixed effects. Since Figure 3 shows evidence of a quadratic relationship between benefit level and likelihood of being affected by the increase, we also explore robustness to excluding claimants with benefit levels above \$1,500 or above \$1,000.

Although we primarily focus on the larger 2002 policy change, we provide additional evidence of a wait time effect by examining a smaller fee ceiling increase that occurs in 2009. Importantly, there are some limitations to studying the 2009 policy. First, any long-lasting dynamic effects of the 2002 policy change may confound estimation of the effects of the 2009 policy. Second, the 2009 change was at the height of the Great Recession and near two other changes in SSDI policy that occurred in 2010.[25] Our empirical strategy can account for secular changes in wait times due to changes in applications, appeals, and awards (Maestas et al., 2015), but only if those changes impact our treated and control groups in the same way. The data suggest these concerns are minimal. Ultimately, we view the 2002 increase as the cleaner natural experiment with a starker treatment, but we see our results on this distinct change in 2009 as further validation of the main results.

Although more-affected and less-affected claimants are similar on many dimensions, there are significant differences (Appendix Table A1). More-affected claimants are slightly older, more educated, and more female. These level differences are not inherently problematic for identification, as long as the parallel trends assumption holds. For this reason we also estimate an event study specification which allows us to examine pre-trends prior to 2002

[25]In 2010, SSA extended direct payment of fees to non-attorney representatives. Between 2010 and 2011, SSA undertook various reporting and monitoring measures to standardize case outcomes across judges at the appeal level.

and visualize the dynamic response to the 2002 (and subsequent 2009) fee cap increase, as follows:

$$Outcome_{it} = \alpha_0 + \phi_t + \sum_{\tau=1997}^{2013} \beta_\tau \mathit{More\text{-}Affected}_i * \phi_\tau + X_i\Gamma + \varepsilon_{it} \tag{4}$$

$Outcome_{it}$, $\mathit{More\text{-}Affected}_i$, and X_i are defined as they are above. ϕ_τ is a vector of year fixed effects from $\tau = 1997$ to 2013. β_τ then represents the average change in $Outcome_{it}$ for more-affected claimants associated with applying in year τ from 1997 to 2013, relative to the omitted year, 1996, and relative to the less-affected claimants in those years. If representatives are responding to the high return on lengthening cases that is specific to more-affected claimants, we should observe an increase in β_τ for years after 2002 and another increase for years after 2009, whether the outcome is wait time or representative fee. The key assumption with this approach is that more-affected and less-affected claimants would have followed similar trends if not for the fee ceiling increase. One testable implication of that assumption lies in the β_τ coefficients from $\tau = 1997$ to 2009. If those coefficients are near zero from 1997 to 2001, that implies that more-affected and less-affected claimants were on similar trends since 1996 and only diverged once the fee was increased in early 2002. Similarly, if the coefficients from 2002 to 2009 exhibit a trend that is roughly flat, that again implies that more-affected and less-affected claimants were on similar trends prior to the second fee increase in mid-2009.

Even though there is a slope change in the level of bindingness at the $500 benefit level, there is not a discrete benefit-level cutoff of who is affected by the fee ceiling increase. For this reason, we relax the treatment and counterfactual difference-in-differences distinction and look at effects for each $100 benefit level bin separately, as follows:

$$Outcome_{it} = \alpha_0 + \sum_{j=100}^{1500} \beta_j (\mathit{In\ Benefit\ Bin\ j}_i * \mathit{PostIncrease}_t) + \gamma_j \mathit{In\ Benefit\ Bin\ j}_i + \phi_t + X_i\Gamma + \varepsilon_{it} \tag{5}$$

The outcomes are the same as above. The β_j coefficients show how the effect of the fee ceiling increase varies across the benefit-level distribution. The under \$100 monthly entitlement bin is the omitted reference group. This approach allows us to verify that our previous results are not sensitive to our choice of the \$500 threshold, but also allows us to determine if the effects are consistent with strategic behavior on the part of representatives. If representatives are being strategic and have exact control over wait times, we would expect the biggest change in wait time after the fee ceiling increase among cases where there is the most to gain. This implies a tradeoff between longer wait times and higher entitlement levels. To see where this lies in the benefit-level distribution, we take claimants that applied prior to the 2002 policy change and calculate how many additional months worth of past due benefits the case could accrue if the binding fee ceiling was lifted from \$4,000 to \$5,300 (the new 2002 limit). For individuals who did not reach the \$4,000 threshold, this value is zero since the initial fee ceiling was not binding so the change is inframarginal. We then plot the average of this for each \$100 benefit level bin in Panel 2 of Figure 3. For low-benefit claimants the extra wait time (and past due benefits) that arise from the fee ceiling increase is negligible, but rises sharply with benefit levels. The excess past due benefits that lead to increased representative fees peak at the \$500 or \$600 monthly entitlement and then steadily declines as benefit levels rise.

In reality, representatives do not have total control over wait times and cannot perfectly predict how long it will take for a case to resolve. This could influence the magnitude of the effects, but we would expect the pattern of effects from equation (5) to follow a similar shape if there is strategic behavior.

6 Results

6.1 Effects of the 2002 Fee Ceiling Increase

Panel A of Table 1 displays the estimated effects on claimant wait time from equation (3). Column 1 includes controls for claimant demographics while column 2 adds primary diagnosis fixed effects. Turning to our preferred specification in column 2 of Panel A, we find that the fee ceiling increase in 2002 increased wait times by 0.85 months, on average, for more-affected claimants. The estimated coefficient is larger but not statistically distinguishable when we exclude diagnosis fixed effects in column 1.[26] From Panel B of Table 1, there is a corresponding $426 increase in the average hypothetical representative fee. These results suggest that SSDI representatives respond to changes in the fee structure that affect their payment. When the fee ceiling increased in 2002, wait times increase specifically for claimants for whom the representative faces a high return to lengthening the case time.

Figure 4 displays the event study coefficients for wait time and hypothetical representative fees. Figure 4 shows that more-affected and less-affected claimants followed similar trends in wait time from 1996-2001. In 2002, however, when the fee ceiling increases, we see a discrete increase in wait time specifically for more-affected claimants. From 2002-2009, the trend in wait time is relatively flat, except for a divergence in 2006 and 2007. In 2009, when the fee ceiling increases again, we again see a larger increase in wait times for more-affected claimants than less-affected claimants. The estimated effects by year for representative fee follow a similar pattern.

In Appendix Table A3, we estimate the effect of the second fee ceiling increase in 2009. Again, we find the fee ceiling increase lengthened wait times, this time by about half a month, on average, for the more-affected claimants. If we scale the coefficients in Tables 1 and A3 by the amount of the respective fee ceiling increases, we recover per dollar increases that are roughly similar in size.

[26] Appendix Table A2 displays results excluding demographic controls and primary diagnosis fixed effects.

6.2 Heterogeneous Effects by Entitlement Level

We next provide the estimates for the different monthly entitlement levels in Figure 5. Average wait times increase sharply with the benefit level after the 2002 fee ceiling increase, peaking in 2005. For claimants with a \$500-600 monthly entitlement the policy change increases average wait times by about 2.8 months. It then gradually declines, leveling around one additional month of wait time for individuals with entitlement levels above \$1,000.[27] This pattern maps closely into what we observe in Panel 2 of Figure 3, indicative of strategic behavior on parts of representatives increasing wait times after the fee schedule changed. It also provides compelling evidence that this is not driven by unobserved secular trends. After the 2002 fee ceiling increase, wait times increase the most precisely for entitlement groups where representatives have the most to gain.

6.3 Robustness

The main threat to identification is that lower entitlement claimants do not provide a valid counterfactual for higher entitlement claimants since, by definition, less-affected claimants have lower incomes. The parallel pre-trends in Figure 4 partially address this concern since they highlight that despite their differences, these groups follow similar trends in wait time and in hypothetical representative fee prior to 2002.[28] As seen in Columns 3 and 4 of

[27]The estimated effects in Figure 5 are considerably larger than the averages in Panel 2 of Figure 3. We believe this is driven by several things. First, representatives do not have total control over wait times and cannot perfectly predict how long it will take for a case to resolve. As such, they might de-prioritize more cases than just those where the fee ceiling is binding. This would lead to larger average effects. Also, Figure 3 relies on the universe of pre-2002 claimants. Secular, programmatic trends in the number of applications and appeal wait times could mean the exact magnitude under strategic behavior differ for the relevant pool of claimants.

[28]We cite parallel pre-trends because the coefficients are near zero in the pre-policy years and they exhibit no discernible trend. Focusing on the confidence intervals, the estimates are precise enough to reject the null that the post-policy coefficient in 2002 is equal to any given pre-policy coefficient. We formalize this even more by implementing the approach of Rambachan and Roth (2022). We find that our estimate for the increase in wait time in 2002 (the first year post-increase) is robust at the 95 percent level to linear violations of the parallel trends assumption and non-linear violations up to a breakdown value around M=0.14. A value of M=0.14 allows the non-linear differential trend to have a change in slope of 0.14 months, a value that is about 87.5% of the standard error on the 2002 coefficient and that is larger than all pre-period coefficients. See Appendix Figure A1.

Appendix Table A1, the more-affected claimants are more similar when we omit claimants with entitlement levels over $1,500 of $1,000 respectively. In Columns 3 and 4 of Table 1 we exclude these high entitlement claimants to focus on individuals with similar income levels and we still see stark differences in wait times after 2002. The same is true when examining the 2009 policy change (Appendix Table A3).

Since the 2002 fee ceiling increase occurred shortly after the 2001 recession, and it remains possible that the less-affected and more-affected groups are differentially impacted by economic downturns. To test if these results are confounded by recessions, we re-estimate the relationship in equation (3), but control for the annual gender by education by region employment to population ratio (constructed using the CPS ASEC) and for education and region fixed effects. Education is split into five groups (less than high school, high school, some college, a college degree, and education missing) while there are 10 regions that roughly correspond to census regions. As seen in Appendix Table A4, including labor market controls when looking at wait times yields similar point estimates, again implying an additional 0.85 months of wait time. In Appendix Figure A2, we plot region-specific treatment effects against a measure of region-level recession intensity, the peak-to-trough change in the employment to population ratio. We do not find a relationship between how hard a region was hit by a recession and our estimated effect of the fee increase in that region. In fact, the direction of the relationship suggests that we find smaller estimates in regions with stronger downturns, suggesting any confounding from these recessions will bias our estimates downward.

During this time period we also saw a rising backlog at the initial claims and hearing levels potentially due, in part, to the sluggish economy (Government Accountability Office, 2007). Rising backlogs could lead to longer wait times in spite of representatives' efforts or actions. This would be problematic for our strategy if the rise in backlogs was more pronounced for *More-Affected* beneficiaries. However, the empirical patterns are not consistent with these effects being driven by backlogs. First, the rise in backlogs begin gradually in 2000 and continue through 2004. Wait times for the more affected group are flat, relative to the less affected group, through 2002 (the year of the policy change) and then experience a large, discrete jump that persists through the end of the sample, which does not match patterns in system wide backlogs. Also, as seen in Figure 5 the effects are largest in the groups where representatives had the most to gain after the policy change. This does not increase monotonically with claimant benefit levels. For these patterns to be driven by backlogs, it would have to be the case that backlogs increased precisely for the groups where the policy was most impactful, but not for claimants with higher or lower benefit levels. We calculate the growth rate in pending initial claims cases starting in 2000 for each SSA region and estimate equation (3) separately for regions that experienced above median and below median growth in backlogs. However, the estimated effects are similar, suggesting the increase in wait time is not driven by regional backlogs (Appendix Table A5). We also find no relationship between region-level growth in pending cases and region-specific treatment effects (see Appendix Figure A3).[29] In fact, we find increasing wait times even in the few regions that experienced a decline in pending cases over this time period. Summing up, we do not find any evidence in support of the concern that our estimates are driven by a gradual increase in backlogs.

Finally, Panel A of Table 2 shows how binding the pre-increase caps are for the four major diagnosis groups: mood disorder, internal systems, musculoskeletal, and infection. We find that among the four groups, claimants with a diagnosis for a disability related to infection are the least likely to reach the pre-increase caps. This is because those cases are resolved more quickly than the cases for other diagnosis groups. As a result, claimants with entitlement levels above $500 are not substantially more likely to reach the pre-increase cap than claimants with entitlement levels below $500. We view this group as providing a useful placebo test for our empirical design. In Panel B of Table 2, we see a similar pattern of the policy on wait time by primary diagnosis. We find increases in wait times for all

[29]GAO calculates backlogs based on the number of pending cases and an SSA-defined optimal target for pending cases (Government Accountability Office, 2007). The 2007 GAO report notes that this target was 400,000 over the time period that they study. Although we do not observe each region's target, the growth rate in pending cases will capture the growth rate in backlogs.

major diagnosis groups except those claimants with a diagnosis for a disability related to infection. The increase in wait time is also monotonically increasing in groups where the policy was more binding. This lends further support to the validity of our difference-in-differences strategy and suggests the estimated increase in wait times is not simply driven by differential trends by income.[30]

6.4 Potential Mechanisms: Effort vs. Timeliness

We next examine whether the increase in wait times could be due to representatives putting more effort into cases rather than a de-emphasis of timeliness. If the fee ceiling increase caused representatives to exert more effort, enter the market, or represent more/different claimants, this could change the number or composition of successful cases, potentially explaining the observed change in wait time. Unfortunately, we cannot directly test how the fee ceiling increase affects the probability of winning using the DAF PUF (since we only observe successful claims), but we can explore this using the Current Population Survey Annual Social and Economic Supplements (CPS ASEC). Using income information in the CPS ASEC to identify "more affected" and "less affected" potential claimants, we estimate how the 2002 fee cap increase affected the probability of receiving SSA benefits among 18 to 60-year-olds with a work inhibiting disability (Appendix Figure A4). There is no evidence that the 2002 policy change affected the rate of benefit receipt, suggesting the fee cap did not induce extensive margin changes in successful cases.

Although we cannot examine the extensive margin using the DAF PUF, we can see how the number or composition of cases changes after the policy changes. We look to see if there is a relative increase in the number of more-affected cases relative to less-affected cases, which would be consistent with a higher success rate and increased effort. As seen in Appendix Figure A5, this does not seem to be the case. If representatives were putting in more effort or if higher quality representatives were entering the market because of the

[30] This pattern is similar across primary diagnosis around the 2009 policy (Appendix Table A6).

fee cap increase, we would expect to see an increase in the number of awards in the more treated group. Based on the event study evidence in Figure 4, we would also expect the selection to be changing discretely in 2002, to explain the discrete increase in wait times. There is not a discrete increase or slope change in the number of more-affected cases relative to the number of less-affected cases around the 2002 fee ceiling increase.[31] This is perhaps a weak test as it combines both changes in take-up and award rates. However, as seen in Appendix Figure A6, there are not discrete changes in successful claimant characteristics after fee ceiling changes which we might expect if representatives increase effort and change the probability of disability award. Instead, average characteristics seem to trend over the entire sample period.

As an additional marker of effort, we explore outcomes related to onset date backdating. After the fee ceiling increase, representatives could increase past due benefits and fee payouts by providing more evidence that the disability began earlier, resulting in a backdated onset date. This would be consistent with an increase in effort, but have more ambiguous welfare implications.[32] However, as seen in Appendix Figure A7, the propensity of early onset dates does not seem to respond to the policy change, but trends in the opposite direction.[33] Since most cases have onset dates prior to the application data, we also explore the average number of months between onset and application in Appendix Figure A8, to see if representatives successfully argue for earlier onset dates. We do not find evidence that representatives are successfully backdating onset dates more for the "more affected" claimants after 2002. We do not find evidence of extensive margin responses to the policy or that representatives are increasing wait times by increasing effort.

Notice that the fee structure implies a specific month when the marginal return to the

[31] The gap does appear to widen in 2009, but once again the increase in applications, appeals, and awards from the concurrent Great Recession make this harder to interpret.

[32] It is worth noting this type of behavior would be beneficial to the claimant as well as the representative as it does not increase the time from application to award but it does increase the months of past due benefits. However, even if we did observe changes here it would suggest that representatives were not acting in their clients' best interest before the change in the fee schedule by leaving benefits on the table.

[33] This is consistent with the early onset rate of low-entitlement claimants catching up to early onset date of high-entitlement claimants.

representative drops to zero. This creates a kink in the relationship between wait time and representative fee that is unique to each monthly entitlement level. We explore potential bunching or heaping in the distribution of wait time around these individual-specific kinks but for brevity, we include this in Appendix B. As noted above, some aspects of the application and appeal process are out of the representative's control. Consistent with this, we observe shifting in the distribution of wait times towards the refund maximizing kink, but it is not precise bunching. This supports the difference-in-difference result suggesting the fee ceiling increase led to a small increase in claimant wait times.[34]

7 Discussion

We estimate that the fee ceiling increase in 2002 is associated with a 0.852 month increase in average wait times for cases where the fee cap is more likely to bind, but is this estimate plausible or economically meaningful? We do not observe which claimants have representation, so this represents an intent to treat effect. To estimate treatment on the treated effects, we must first identify the "effective" treated group. This will be composed of individuals that have representation and for whom the policy change potentially affects the representative's incentives. Hoynes et al. (2021) report that 15 percent of initial claims had legal representation in 2014. However, this does not include cases where representation was brought on in the appeal stage which is closer to 80 percent. If we take the ratio of the annual number of award payments made to representatives (to proxy for the number of claims with representation), divided by the annual number of worker awards, we estimate higher representation rates of 29.4 percent in 2002 with an average of 32.4 percent between 2002 and 2009.

[34]As a final piece of evidence supporting strategic behavior we explore differences by geographic market power. If there are fewer disability representatives in an area and less competition, representatives might have more market power and a greater ability to extract a higher total fee by de-emphasizing timeliness without fear of competition capturing the market. Using the appellate market share maps provided by Hoynes et al. (2021), we estimate the average of market share across the nine geographic regions, and find that the increase in wait times after the 2002 fee ceiling increase is larger in places with more concentrated market shares.

Figure 3 suggests that for approximately 10 percent of claimants, the 2002 fee cap increase from \$4,000 to \$5,300 relaxed the maximum fee constraint. However, this is an under-estimate of the total share of claimants who are "treated" by the fee increase. There are many individuals for whom the initial policy was not binding, but the new policy still increases the amount the representative can receive. For example, consider an individual in 2001 for whom the representative fee was \$3,999. The \$4,000 cap is not technically binding in this case, but the change to a \$5,300 cap significantly increases the amount of revenue associated with extra wait time. In Appendix Table A7, we document that in approximately 19% of cases, representatives would benefit more from a 4 month wait under the new cap than the old cap. This share jumps to 24% when considering the return to 6 months of additional wait time, and 48% when considering the return to 12 additional months. These shares are likely still under-estimates of the effective treated group. As we note above, there are many determinants of wait time that are outside of the claimant or representative's control so neither claimants nor representatives know how long it will take for their case to resolve. As such, representatives cannot only apply delays to the "binding" claims, because they will not know if it is binding or not until the case is resolved. Representatives can make predictions based on observable characteristics, but they cannot know definitively for a given case if extra wait time will be marginal or inframarginal.

If we use the overall level of representation (32.4 percent) and implicitly assume that all represented cases are to some extent treated, we estimate an average treatment effect among the represented of 2.6 months (0.852/0.324).[35] If we scale by the "effective" treated group – say the 19% of cases in which the representative would benefit more from a 4 month wait under the new cap than the old cap – then we also want to scale the coefficient by the representative share among *those* cases, not *all* cases. Cases that reach this level of wait time are cases that have gone to appeal, and approximately 80% of appeal cases have representation. Thus, our estimate of the average treatment effect on the effective treatment

[35]This is assuming the policy changes do not affect representation at the extensive margin, consistent with evidence in Appendix Figures A4 – A6.

group is 5.6 (0.852/(0.8*0.19)) months. Based on an average wait time in our sample of 11.6 months, this would represent a 22 percent or 48 percent respective increase in wait time for claimants.

As noted above, after each stage of SSA evaluation, claimants and representatives have 60 days to resubmit documentation for consideration, meaning for many treated cases representatives potentially have four months or more of discretion. Both of these point estimates are not statistically distinguishable from the amount of delay consistent with the 60 day re-submission deadlines between SSA actions. This does not account for additional delays from failure to submit complete documentation and evidence, as has been alleged by some field office and DDS employees (Social Security Advisory Board, 2012).

Even though longer wait times do not reduce the total income transfer to claimants, an increase in wait time could still harm claimant outcomes.[36] To be eligible for disability insurance, claimants must maintain earnings below the substantial gainful activity threshold throughout the application process and after award. As Deshpande et al. (2021) note, gaining access to disability insurance significantly reduces the likelihood a claimant sells their home, forecloses, or declares bankruptcy. This would suggest applicants are liquidity constrained and cannot borrow against their future income to avoid negative financial events. If they are liquidity constrained, slightly higher past due benefits in several months might not be as valuable as a smaller amount of past due benefits now. Longer wait times also increase the risk that claimants do not live to receive awarded benefits. Between 2008 and 2019, nearly 110,000 disability insurance applicants died while awaiting their final decision (Government Accountability Office, 2020). Longer wait times can also affect those claimants that are not awarded disability insurance. Since claimants cannot work while awaiting a decision, claimants who wait longer experience more human capital decay and report lower future earnings (Autor et al., 2017).

These results are consistent with work by Hoynes et al. (2021), who find that retaining

[36] Longer wait times might result in lower total lifetime benefits if 25% of past due benefits still falls below the fee ceiling as a larger share of lifetime benefits is transferred to the representative.

representation significantly increases the initial approval rate, but has no effect on the overall approval rate. They also find that retaining representation is associated with a statistically imprecise, but large, 9 day increase in field office processing time (143% of the mean), a statistically imprecise, but large, 63 day increase in DDS processing time (68% of the mean) and a statistically significant 31 percentage point increase in the probability of onset to DDS decision taking longer than five months.[37] Despite these patterns, retaining representation significantly decreases total processing time on average by 316 days, largely by dissuading claimants with weaker cases from appealing SSA decisions. We once again underscore that Hoynes et al. (2021) explore a distinct question: does representation improve the odds of winning a disability award and change the application process? Representatives can both increase the probability of award and respond to fee structure incentives. Among successful claimants, representation might increase claimant wait times. We provide complementary evidence that, although retaining a representative might improve some outcomes, the existing fee structure might lead to a smaller net benefit than alternative fee structures.

Even though raising the fee ceiling results in longer wait times, we are hesitant to make claims about overall welfare or optimal policy structure. Changing the fee cap could have long-run dynamic effects. Over time, disability representatives might change the way they select clients potentially leaving claimants with low benefits or complex cases without viable representation options. On the other hand, increasing fee caps could make disability law more lucrative, encouraging more lawyers to enter the field and increase competition over the long term. As suggested in footnote 34, an increase in competition could partially mitigate the effect of the fee cap on wait times in the long-run if the higher fee rates encourage new representatives to enter the market. Unfortunately, these are phenomena we are unable to explore with the current data. Recent discussions have proposed alternative fee structures, including minimum fees, flat fees, reverse time-dependent fees, and various actuarially fair schemes that depend on case characteristics (Abt Associates, 2019). Many

[37] As the authors note, this could be driven by representatives pushing for an earlier onset date, which increases past due benefits but does not affect the actual number of months the claimant waits.

of these designs eliminate or reduce the incentive effects documented here, but potentially induce representatives to be more selective in the claimant/representative matching process or solicit claimants who may benefit little from legal aid. It is unclear whether these changes would improve total welfare.

Instead of changing the fee structure, changing what counts as past due benefits for representative compensation could reduce incentives to prolong wait times. Although there are many stages in the application process that are out of representatives' hands, they do have near complete control over the time that elapses between denial and re-application or denial and appeal (the 60 day deadline) once the claimant seeks representation. If this time is not counted when calculating representative fees, the incentive to delay or de-prioritize cases is largely removed. This sort of change could correct for the fee structure incentives without implementing a new fee regime with uncertain equilibrium outcomes.

8 Conclusion

The SSDI application and appeal process is often drawn out and complicated. This leads many claimants to seek help from legal representation. The compensation rules of SSDI representation are dictated by the Social Security Administration to protect SSDI claimants. However, the structure of these rules may also create incentives for SSDI representatives to only take on cases that are likely to succeed and then to allow cases to drag on for long periods of time to increase past due benefits and representative compensation. In this paper, we use difference-in-differences to provide suggestive evidence that the structure of representative compensation rules leads to longer wait times for claimants. We find that the maximum representative fee increase in 2002 led to a reduced form average increase in wait time of 0.85 months, or between 2.6 and 5.6 months once we scale by the representation rate.

Data limitations in the DAF PUF preclude us from answering interesting questions that could be examined using richer data. Data on individual claimant representation would lend itself to alternative identification strategies. Detailed information on representation agreement entry date could also be used to examine high frequency variation in close proximity to the fee ceiling changes. Information on the actual representatives could also reveal whether wait time effects are concentrated among certain representative types or groups or are broadly observed. Richer data could further shed light on how representative fees affect who retains representation or how this affects the probability of getting an award. Without clear information on the application and appeal process in the DAF PUF, we also cannot examine other behaviors, like representatives delaying or withholding documentation to prolong the process. Unfortunately, much of the administrative data that contains representative information does not go back far enough to examine the effect of maximum representative fee increases (Hoynes et al., 2021). Despite these data limitations, we still provide the first evidence of how the representative fee structure affects claimant wait times.

This work sheds new light on how the incentives of social program gatekeepers can affect applicant outcomes. Many disability insurance claimants rely on representatives' expertise to successfully navigate and participate in the program. Current SSA regulation of representative compensation creates incentives to delay and results in longer wait times for claimants. These results also apply to other social programs and legal representative settings. Many programs rely on administrators and intermediaries. Policy architects should consider their incentives and objectives when designing social programs.

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Tables and Figures

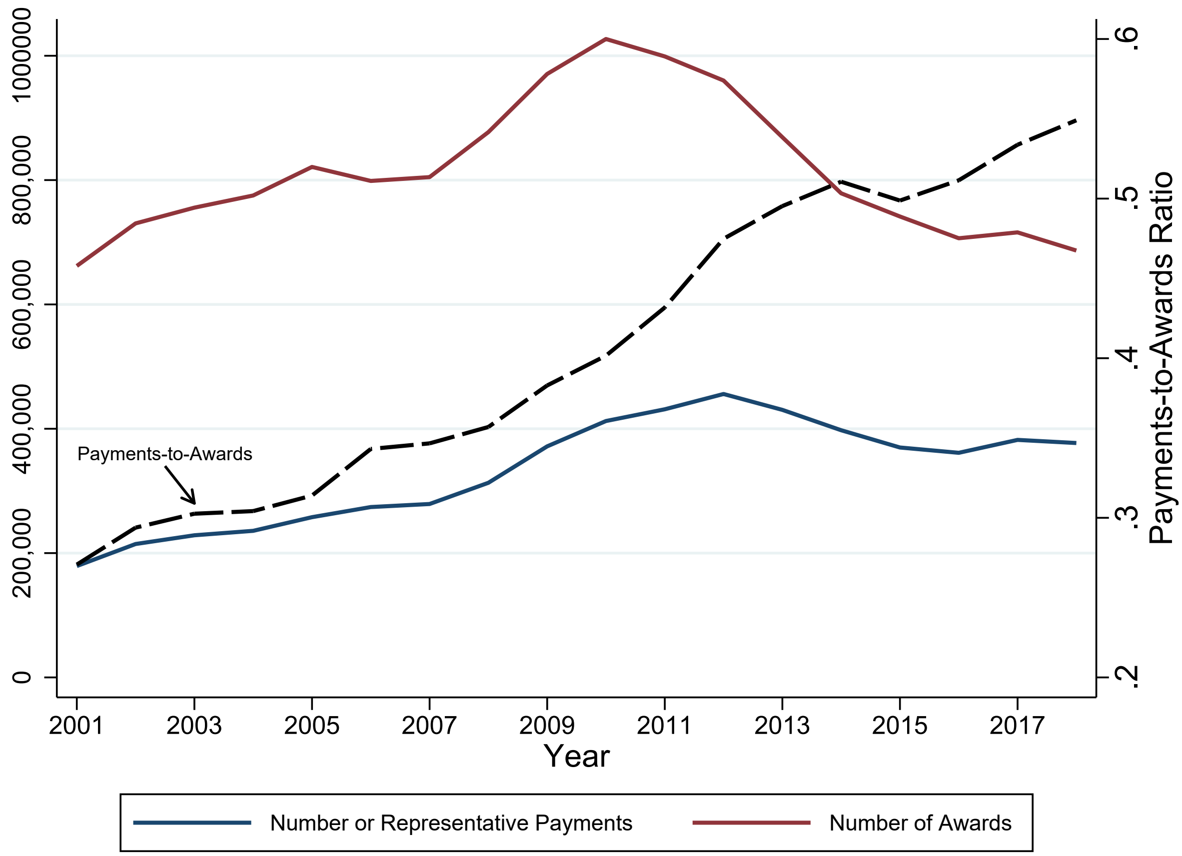

Figure 1: Number of Worker Awards and Number of Payments Made to Claimant Representatives by Year

Notes: The figure above plots the total number of awards to workers (dependent awards are excluded) and the number of fee payments made to claimant representatives by year. The total amount paid has also grown over time from $40 million per month in 2000 to approximately $100 million per month in 2019. The total amount paid per fee agreement, on the other hand, rose substantially from 2000 to 2007, increasing from $2,250 to $3,500, but then fell from 2007 to 2014, bottoming out at around $2,800. As of 2019, the average amount paid per fee agreement is about $3,000.

Source: Authors' own calculations from SSA's "Statistics on Title II Direct Payments to Claimant Representatives" and the "Annual Statistical Report on the Social Security Disability Insurance Program".

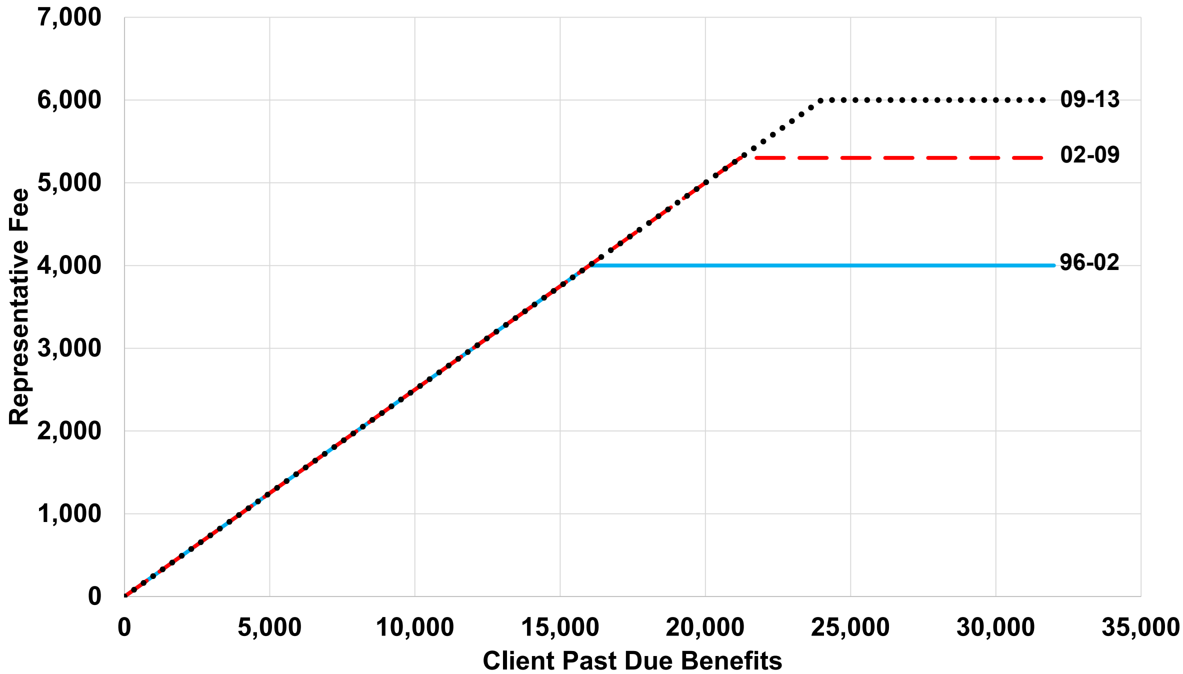

Figure 2: Representative Fee Schedule Over Time

Notes: Representative compensation schedule plotted over time. This structure was first adopted by the SSA in 1990, and a fee ceiling was set at $4,000. The fee ceiling was increased to $5,300 on February 1, 2002 and to $6,000 on June 22, 2009. Focusing on our years of interest: the schedule in effect from 1996-2002 is plotted with the solid blue line, the schedule in effect from 2002-2009 is plotted in the dashed red line, and the schedule in effect from 2009-2013 is plotted in the dotted black line. The fee ceiling applies to agreements made between the claimant and representative after that date, regardless of their initial application date.

Source: Authors' own calculations.

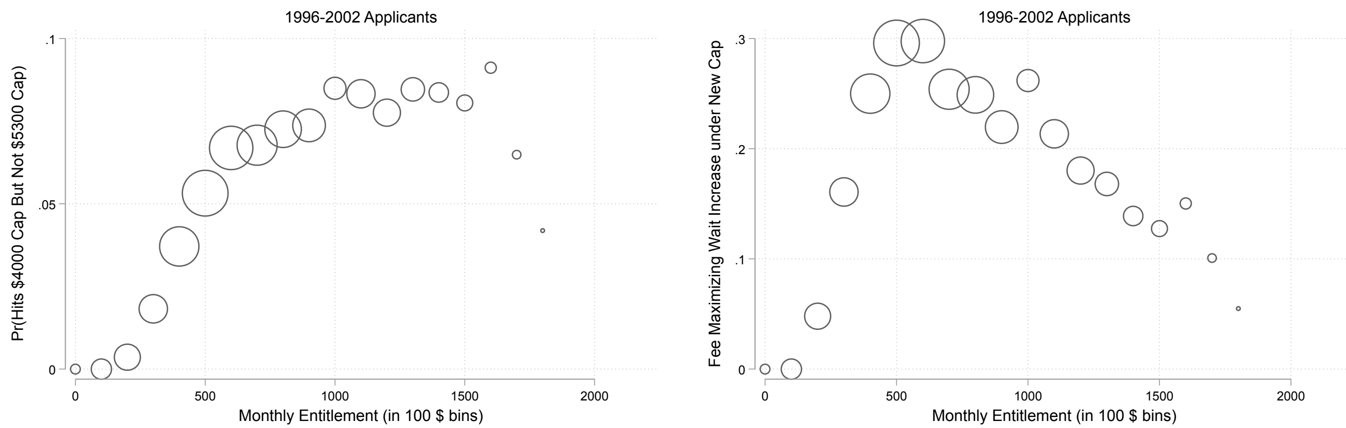

Figure 3: Predicted Impact of 2002 Fee Ceiling Increase on Bindingness and Strategic Wait Time Changes

Notes: Sample restricted to claimants who applied between January 1, 1996 to February 1, 2002, prior to the fee ceiling increase. The left panel shows how much less binding the future fee ceiling would be by claimant entitlement level. Within 100 dollar monthly entitlement bins we calculate the share of claimants for whom 25 percent of past due benefits exceeds the prevailing fee ceiling but would not have exceeded the new fee ceiling. The right panel shows the average change in wait time after the fee ceiling increase that would yield increased revenue for the representative. Within 100 dollar monthly entitlement bins we calculate how many additional months worth of past due benefits the case could accrue if the binding fee ceiling was lifted from $4,000 to $5,300 (the new 2002 limit). For individuals who did not reach the $4,000 threshold, this value is zero since the initial fee ceiling was not binding so the change is inframarginal. In other words, this measure captures the average additional months worth of revenue a representative could capture if they had full control of wait time. Wait time is the period between the entitlement date and award date. The fee regulation from the application year is used. If the individual enters an agreement with representation at a later day, more generous compensation parameters might apply. The patterns looks similar if we account for claimant gender, age, and primary diagnosis.

Source: Authors' own calculations from the SSA DAF-PUF.

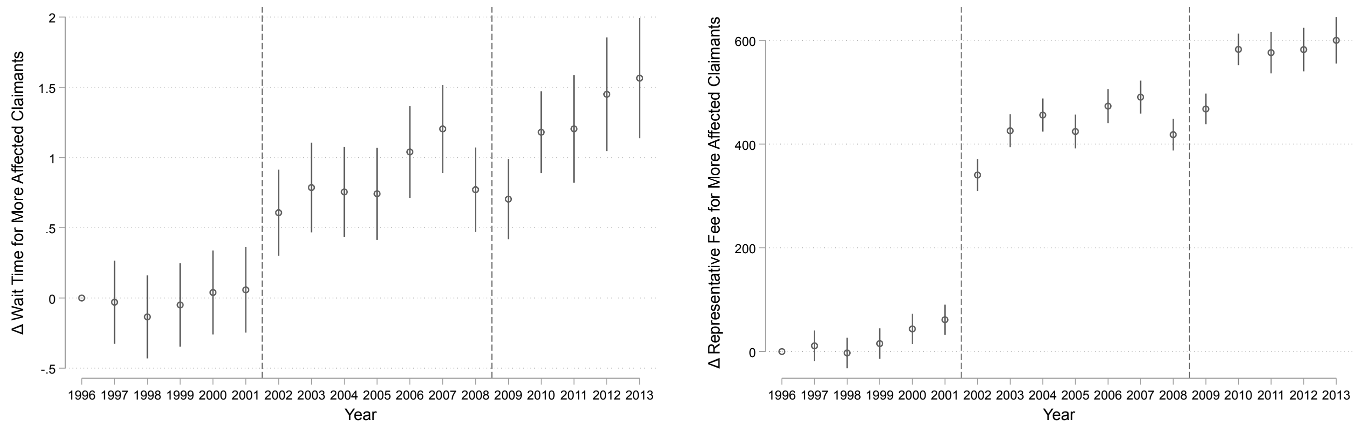

Figure 4: Change in Wait Time and Representative Fees For High Benefit Claimants Relative to Low Benefit Claimants Since 1996

Notes: Sample restricted to claimants who applied between 1996 and 2013. Coefficients estimated from equation (4). Each coefficient represents the change in wait time –entitlement date to approval date– in months (left panel) or the hypothetical representative fee in dollars (right panel) since 1996 for claimants with above $500 in monthly entitlement relative to claimants with below $500 in monthly entitlement. Sex, age at application, primary diagnostic code fixed effects, $100 monthly entitlement bin fixed effects, and year of application fixed effects are included. We include a dashed line after 2001 because the 2002 change occurs in February 2002, and we include a dashed line after 2008 because the 2009 change occurs in June 2009. 95 percent confidence intervals based on robust standard errors are shown.

Source: Authors' own calculations from the SSA DAF-PUF.

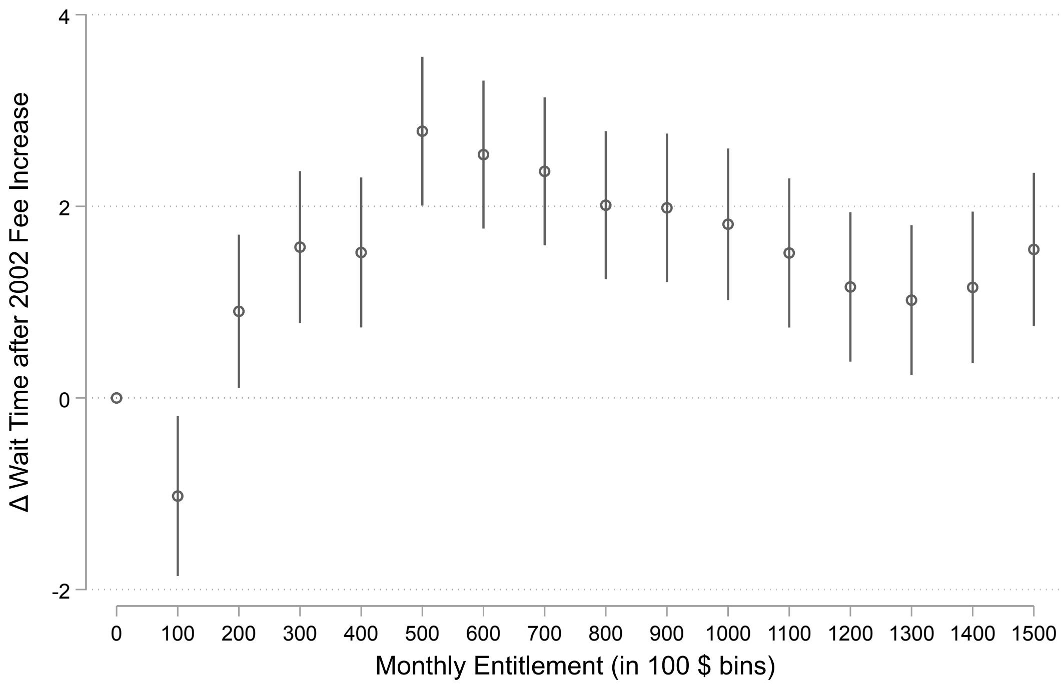

Figure 5: Heterogeneous Effect of the 2002 Fee Ceiling Increase on Wait Time by Monthly Entitlement Level

Notes: Sample restricted to claimants who applied between 1996 and June 2009. In this sample, 17.5% of claimants are in the $0-$500 monthly benefit range. Coefficients estimated from equation (5). Each coefficient represents the change in wait time –entitlement date to approval date– in months after the 2002 policy change by monthly entitlement (in $100 bins). The under $100 bin is the omitted reference group. Sex, age at application, primary diagnostic code fixed effects, $100 monthly entitlement bin fixed effects, and year of application fixed effects are included. 95 percent confidence intervals based on robust standard errors are shown.

Source: Authors' own calculations from the SSA DAF-PUF.

Table 1. Effect of 2002 Fee Ceiling Increase on Claimant Wait Time and Hypothetical Representative Fee

	(1)	(2)	(3)	(4)
Panel A. Wait Time from Entitlement to Approval (in Months), 1996-2009				
More Affected x Post-2002	1.116***	0.852***	0.864***	1.163***
	(0.0649)	(0.0622)	(0.0624)	(0.0647)
Constant	14.63***	7.224***	6.907***	6.634***
	(0.0737)	(0.440)	(0.461)	(0.705)
Observations	1,124,297	1,124,297	1,036,503	786,427
R-squared	0.057	0.149	0.144	0.130
Panel B. Hypothetical Representative Fee (in Dollars), 1996-2009				
More Affected x Post-2002	469.0***	426.4***	427.1***	429.2***
	(6.484)	(6.257)	(6.270)	(6.567)
Constant	2,270***	865.4***	826.7***	901.0***
	(9.660)	(107.1)	(107.0)	(137.6)
Observations	1,124,297	1,124,297	1,036,503	786,427
R-squared	0.092	0.194	0.195	0.193
Primary Diagnosis FEs	NO	YES	YES	YES
Sample Restriction	NO	NO	Excluding $1,500+	Excluding $1,000+

Notes: *** p<0.01, ** p<0.05, * p<0.10. Robust standard errors in parentheses. This table estimates the difference-in-differences model from equation (3) for the 2002 fee ceiling increase. We limit the sample to claimants who applied under the $4,000 cap or the $5,300 cap. We then define Post-2002 as equal to one if the claimants applied under the $5,300 cap and zero otherwise. In Panel A, the outcome variable is likelihood that a claimant reaches the $4,000 cap but would not reach the $5,300 cap for the sample of claimants who applied under the $4,000 cap.the claimant's wait time from entitlement to claim approval (in months). In Panel B, the outcome variable is the hypothetical representative's fee (in $) based on the claimant wait time, claimant benefit level, and the fee ceiling in effect at the time of application. For both panels, column 1 includes controls for claimant sex and age and column 2 includes those controls plus primary diagnosis fixed effects. Column 3 includes the same controls as column 2 but excludes claimants with benefit amounts above $1,500 per month, and column 4 includes those same controls controls but excludes claimants with benefit amounts above $1,000 per month.

Table 2. Effect of 2002 Fee Ceiling Increase by Primary Diagnosis

Primary Diagnosis	Mood Disorder	Int. Systems	Musc. Skel.	Infection
	(1)	(2)	(3)	(4)
Panel A. Likelihood of Reaching Pre-Increase Cap but Not Post-Increase Cap, 1996-2002				
More Affected	0.0506***	0.0421***	0.102***	0.0182***
	(0.00140)	(0.00125)	(0.00223)	(0.00112)
Constant	-0.00467*	0.0403***	0.122***	0.0303***
	(0.00277)	(0.00264)	(0.00608)	(0.00229)
Observations	95,724	119,947	67,963	72,942
R-squared	0.012	0.006	0.020	0.003
Panel B. Wait Time from Entitlement to Approval (in Months), 1996-2009				
More Affected x Post-2002	0.742***	0.617***	0.814***	-0.0916
	(0.114)	(0.124)	(0.146)	(0.150)
Constant	10.90***	12.77***	31.96***	8.219***
	(0.142)	(0.130)	(0.211)	(0.120)
Observations	284,946	339,420	256,576	188,140
R-squared	0.023	0.055	0.087	0.056

Notes: *** p<0.01, ** p<0.05, * p<0.10. Robust standard errors in parentheses. In Panel A, we limit the sample to claimants who applied under the $4,000 cap. We then define More Affected as equal to one if the claimant's monthly benefits exceed $500 and zero if not. The outcome variable is an indicator that a claimant reaches the $4,000 cap but would not have reached the $5,300 cap. Panel B estimates the difference-in-differences model from equation (3) for the 2002 fee ceiling increase. We limit the sample to claimants who applied under the $4,000 cap or the $5,300 cap. We then define Post-2002 as equal to one if the claimants applied under the $5,300 cap and zero otherwise. The outcome variable is the claimant's wait time from entitlement to claim approval (in months). All columns include demographic controls (sex and age). Columns 1-4 are limited to claimants with diagnoses of: mood disorder, int. systems, musculoskeletal, and infection, respectively.

Online Appendix A. Supplementary Analyses

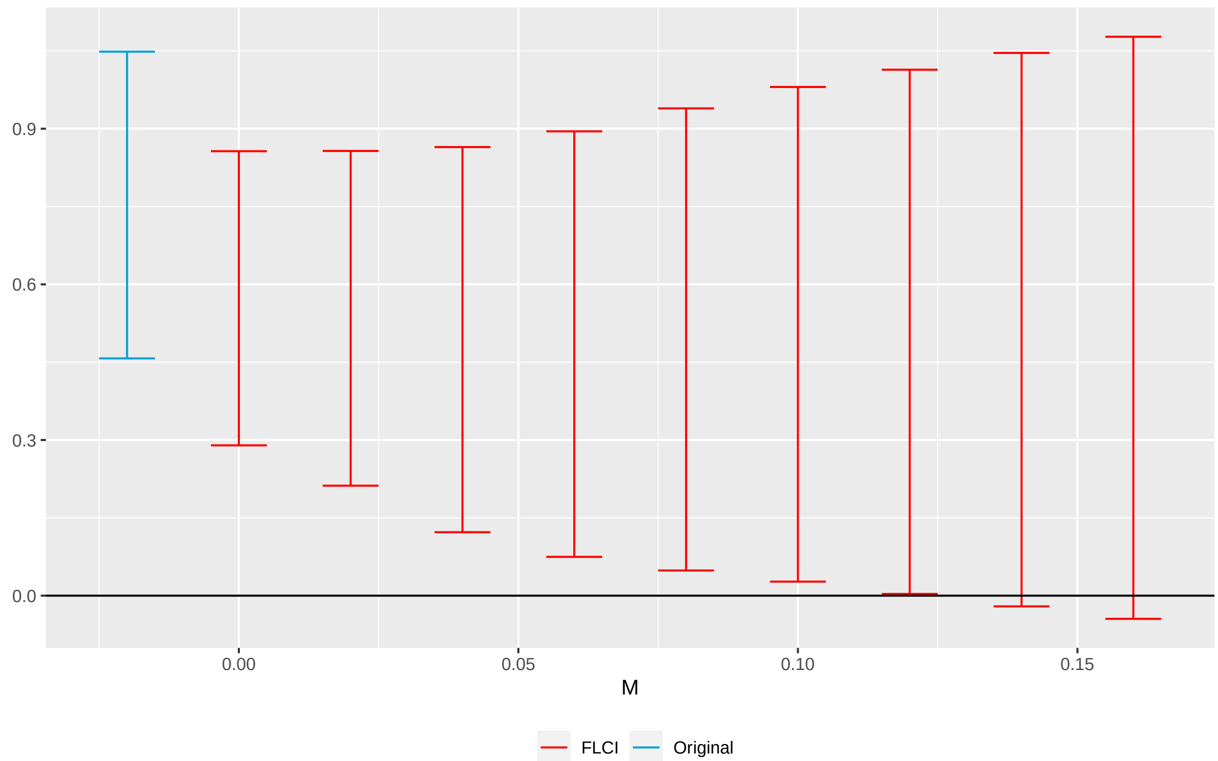

Figure A1: Robustness of Wait Time Estimate to Linear and Non-Linear Trend Differences

Notes: Sample restricted to claimants who applied between 1996 and 2009 (excluding the latter half of 2009 after the second fee increase). Coefficients estimated from equation (4). After estimating the event study, we store the estimate on wait time from the first year post-increase and test robustness of that estimate to linear and non-linear trend differences. Specifically, we implement the approach from Rambachan and Roth (2022). The first set of confidence intervals in this figure shows the main estimate for 2002. The following confidence intervals show robustness to various values of M. A value of M=0 allows for linear violations of the parallel trends assumption. We find that our estimate for the increase in wait time in 2002 is robust at the 95 percent level to linear violations of the parallel trends assumption and non-linear violations up to a breakdown value around M=0.14. A value of M=0.14 allows the non-linear differential trend to have a change in slope of 0.14 months, a value that is about 87.5% of the standard error on the 2002 coefficient and that is larger than all pre-period coefficients.

Source: Authors' own calculations from the SSA DAF-PUF.

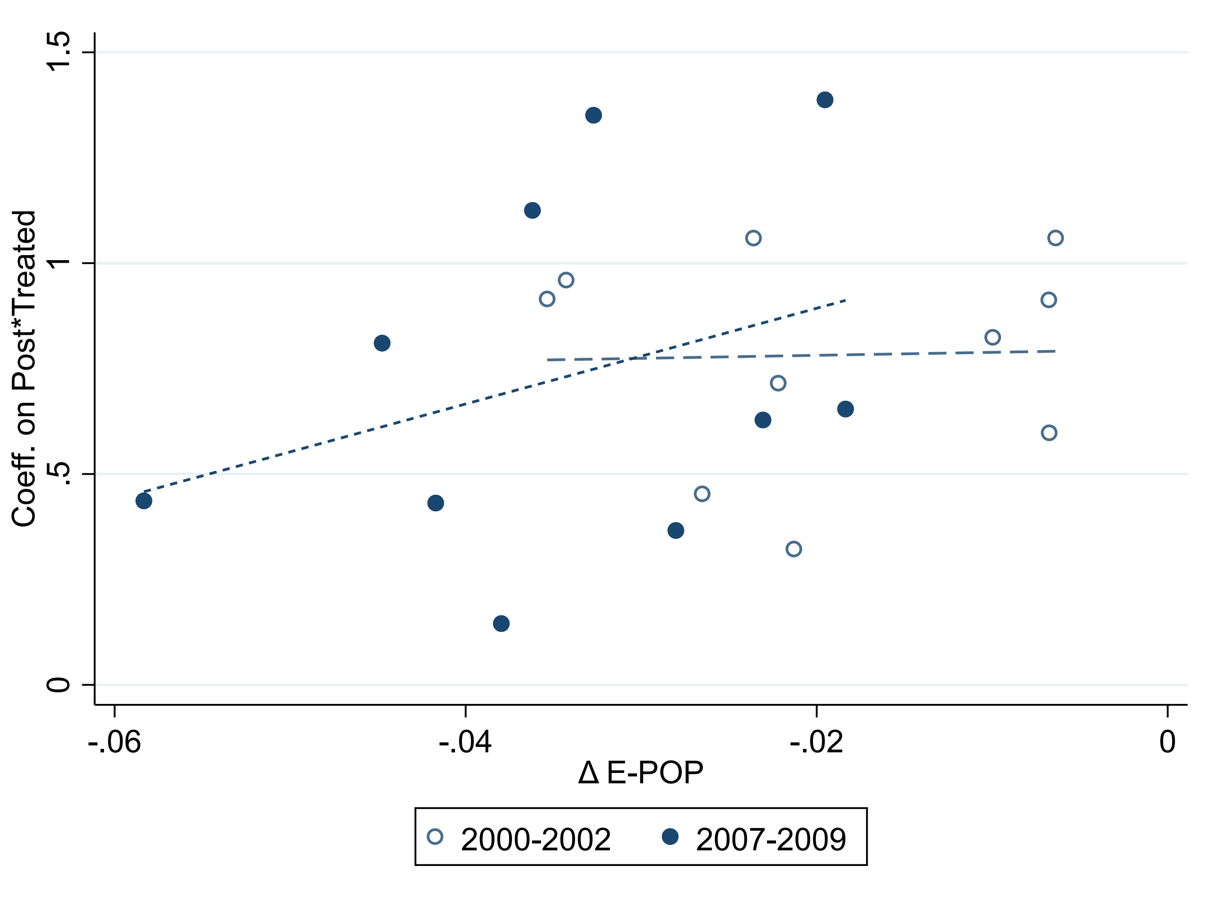

Figure A2: Relationship between Region Estimates and Economic Conditions

Notes: Sample restricted to claimants who applied between 1996 and 2013. Coefficients estimated by region from equation (3). Each coefficient represents a region-level estimate of the change in wait time after the fee cap increase for claimants with above $500 in monthly benefits relative to claimants with below $500 in monthly benefits. Sex, age at application, primary diagnostic code fixed effects, and $100 monthly entitlement bin fixed effects. These coefficients are plotted against the region-level change in the employment to population ratio from 2000-2002 and 2007-2009. The long dashed line plots the linear fit between the region-level estimates based on the 2002 fee increase and the 2000-2002 change in the employment to population ratio. The short dashed line plots the linear fit between the region-level estimates based on the 2009 fee increase and the 2007-2009 change in the employment to population ratio.

Source: Authors' own calculations from the SSA DAF-PUF.

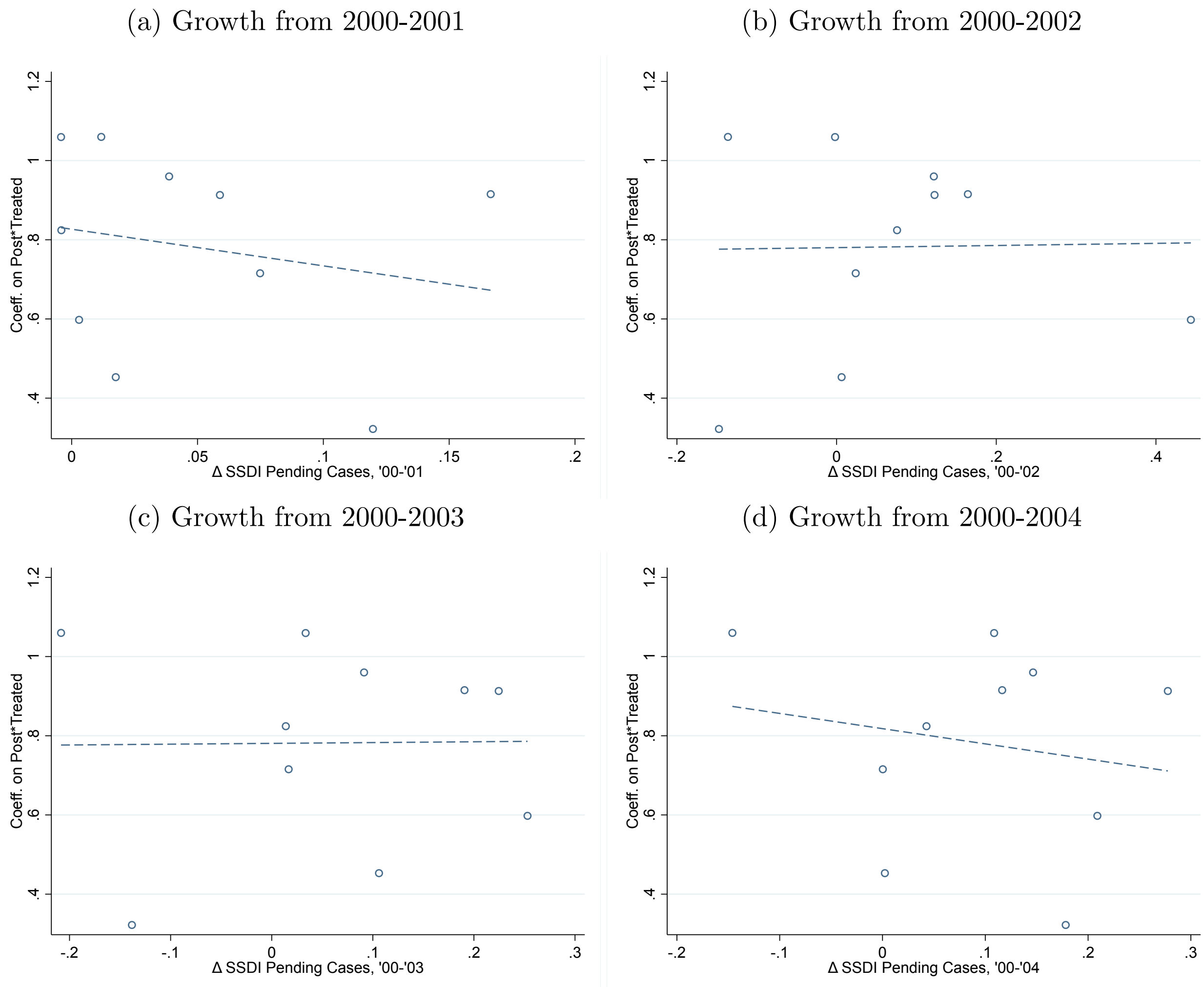

Figure A3: Relationship between Region Estimates and Growth in Pending Cases

Notes: Sample restricted to claimants who applied between 1996 and 2009. Coefficients estimated by region from equation (3). Each coefficient represents a region-level estimate of the change in wait time after the fee cap increase for claimants with above \$500 in monthly benefits relative to claimants with below \$500 in monthly benefits. Sex, age at application, primary diagnostic code fixed effects, and \$100 monthly entitlement bin fixed effects. These coefficients are plotted against the region-level change in pending initial claims cases from 2000-01, 2000-02, 2000-03, and 2000-04. We show results for various definitions of growth in backlogs without taking a stance on which time period is most relevant. The long dashed line plots the linear fit between the region-level estimates based on the 2002 fee increase and the change in pending initial claims cases. The growth rate in pending cases is based on SSDI cases only, although we find similar results when using the growth rate in pending cases for all disability benefits.

Source: Authors' calculations from the SSA DAF-PUF and SSA State Agency Monthly Workload Data.

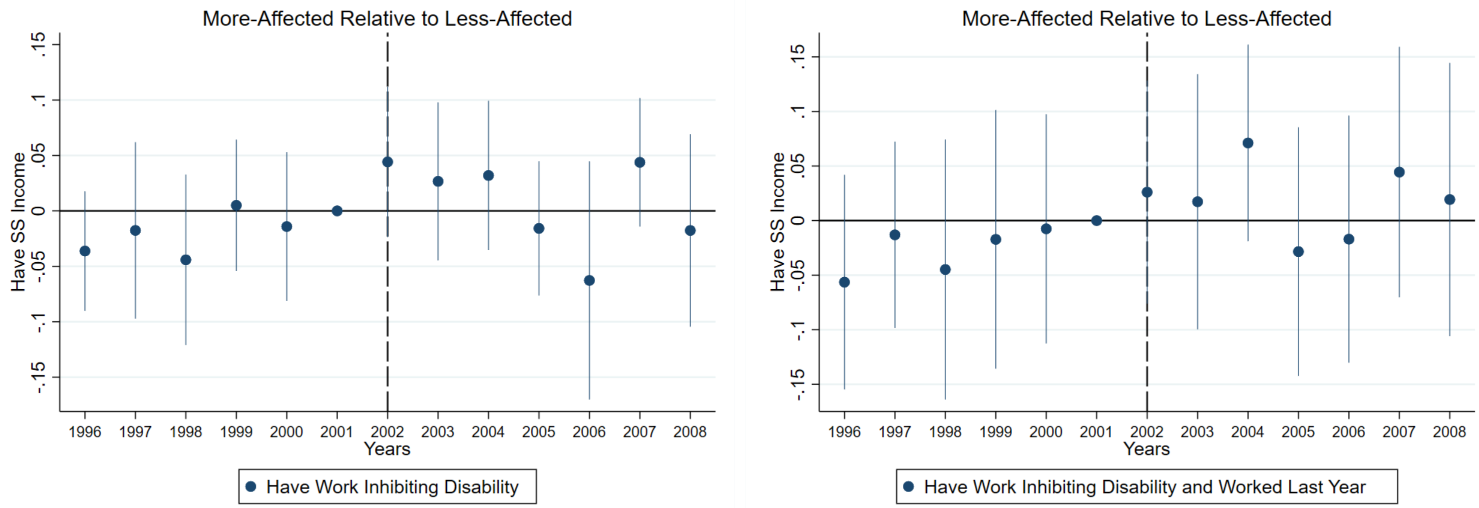

Figure A4: Extensive Margin Responses: Event Study Evidence from SSA Income Receipt in the CPS ASEC

Notes: Sample restricted to 1996-2008 CPS ASEC respondents between the ages of 18 and 60 who reported having a disability that limits or prevents work in Panel A. The sample is further restricted in Panel B to only include individuals who had wage income in the previous 12 months, to capture flows onto SSDI. The outcome is a binary measure that equals one if the individual received income from the Social Security Administration (SSA). Among working age individuals (18-60) with a work inhibiting disability in the 2001-2022 CPS ASEC, approximately 92 percent of individuals receiving income from SSA receive payments for disability insurance. Using wage income in the previous year, we estimate the relationship between inflation indexed wage earnings and observable characteristics (race/ethnicity, gender, education level, age fixed effects, year fixed effects, state fixed effects, metropolitan area fixed effects, and marital status) among individuals with non-zero wage earnings. We then predict each individual's annual wage earnings based on their observable characteristics, regardless of if they had actual wage earnings or not. With this predicted income measure, we then construct a predicted AIME, predicted PIA, and a binary measure corresponding to "More Affected", as defined throughout the manuscript. Sample is weighted using CPS ASEC weights. Fixed effects for state and year are included. Standard errors are corrected for clustering at the state-level with 95 percent confidence intervals reported.

Source: Authors' own calculations from the CPS ASEC.

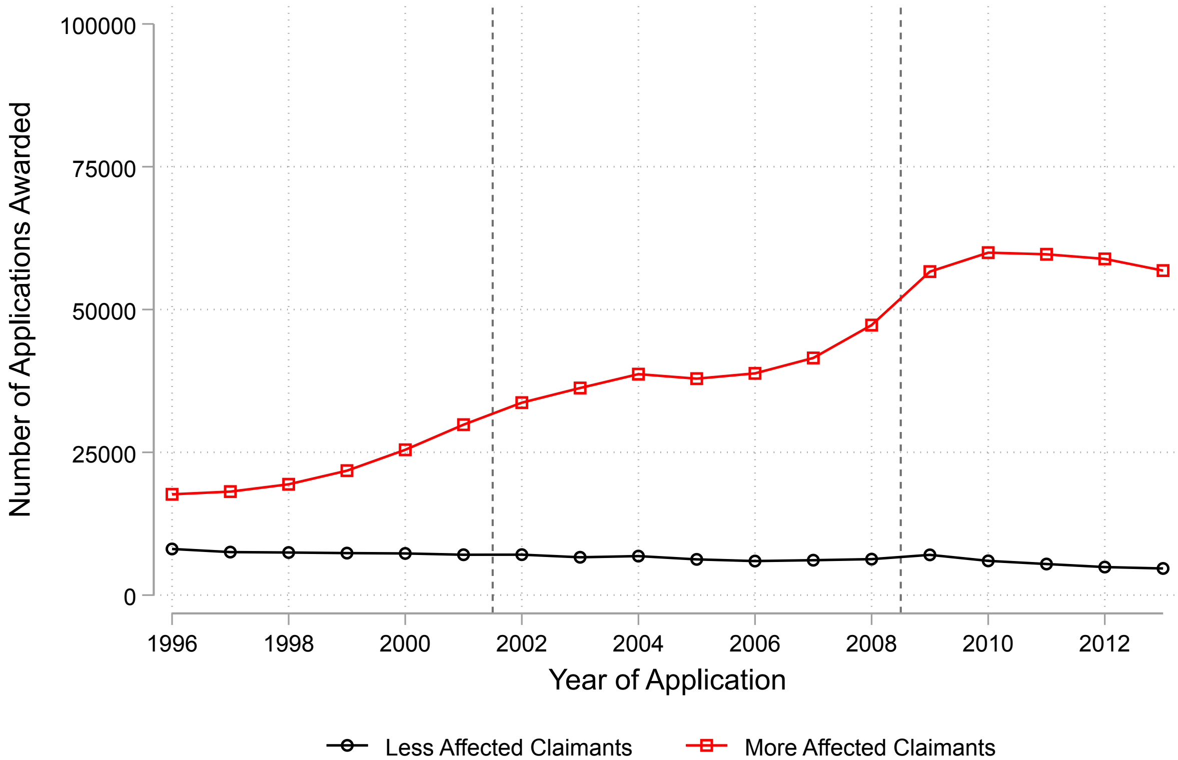

Figure A5: Number of Less Affected and More Affected Awards by Application Year

Notes: Sample restricted to claimants from the 2018 extract who applied between 1996 and 2013. More affected claimants had monthly entitlements above $500 while less affected claimants had entitlements below $500.

Source: Authors' own calculations from the SSA DAF-PUF.

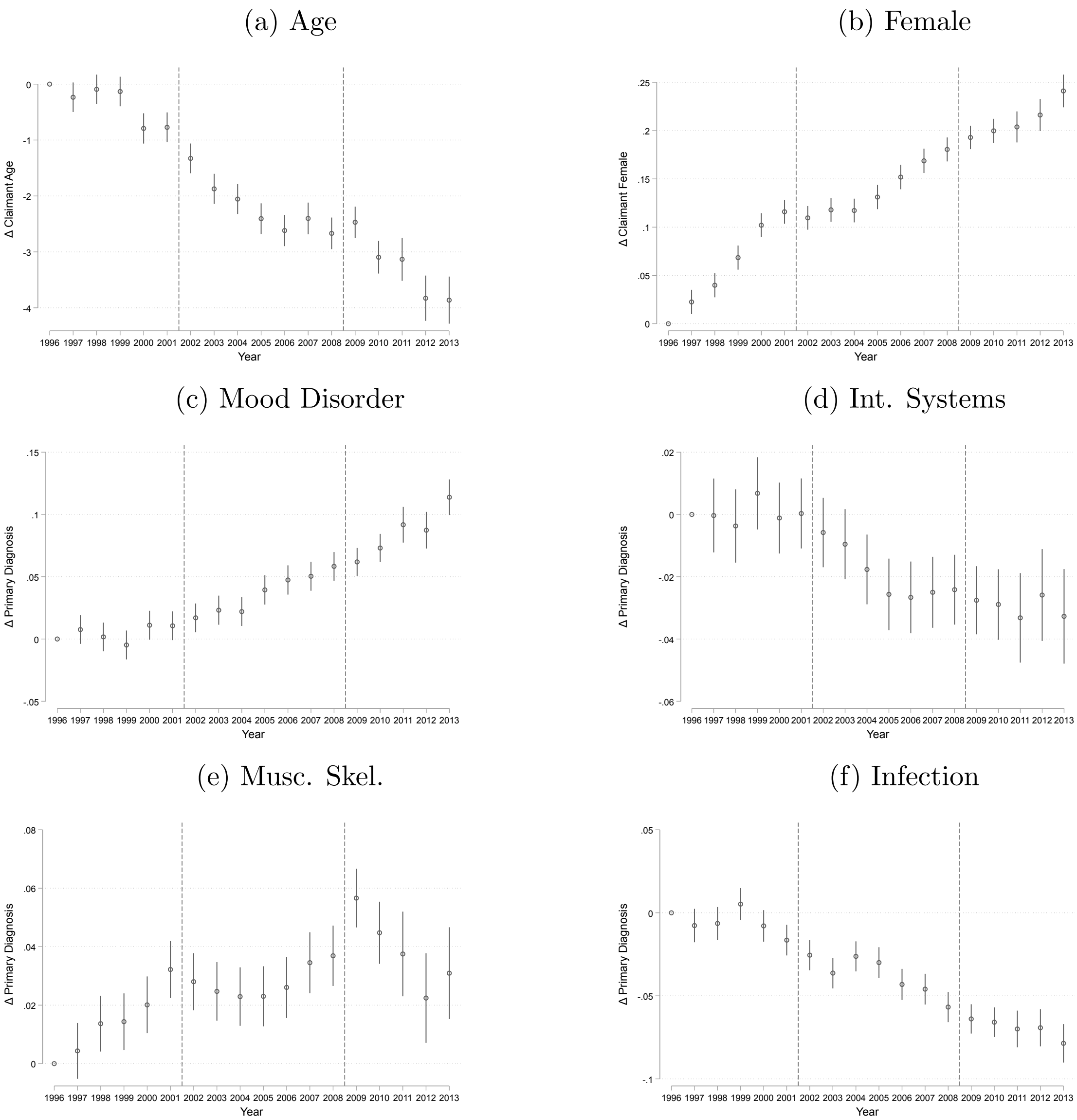

Figure A6: Claimant Characteristics for High Benefit vs. Low Benefit Claimants since 1996

Notes: Claimant data from the DAF PUF. Sample restricted to primary claimants that only have one entitlement date and applied in 1996 or later. The coefficients represent the change in the share of more-affected claimants with the given characteristic relative to less-affected claimants since 1996. Year of application fixed effects and \$100 monthly entitlement bin fixed effects are included.

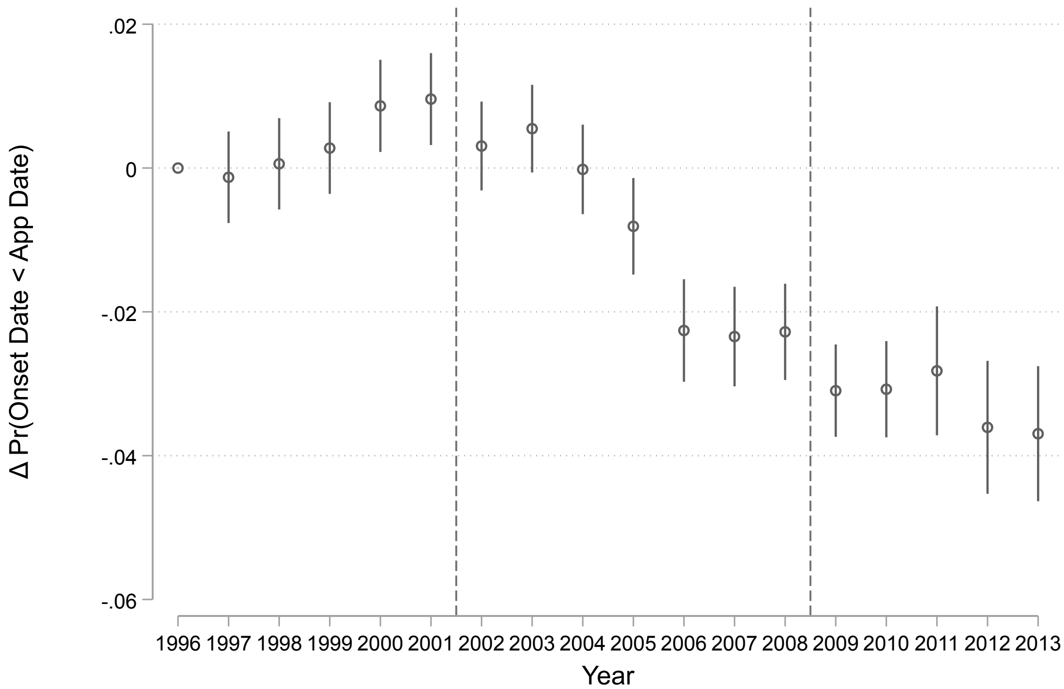

Figure A7: Change in Having Onset Date Before Application Date For High Benefit Claimants Relative to Low Benefit Claimants Since 1996

Notes: Sample restricted to claimants who applied between 1996 and 2013. Coefficients estimated from equation (4). Each coefficient represents the change in likelihood of having an onset date prior to application date since 1996 for claimants with above $500 in monthly benefits relative to claimants with below $500 in monthly benefits. Sex, age at application, primary diagnostic code fixed effects, $100 monthly entitlement bin fixed effects, and year of application fixed effects are included. We include a dashed line after 2001 because the 2002 change occurs in February 2002, and we include a dashed line after 2008 because the 2009 change occurs in June 2009. 95 percent confidence intervals based on robust standard errors are shown.

Source: Authors' own calculations from the SSA DAF-PUF.

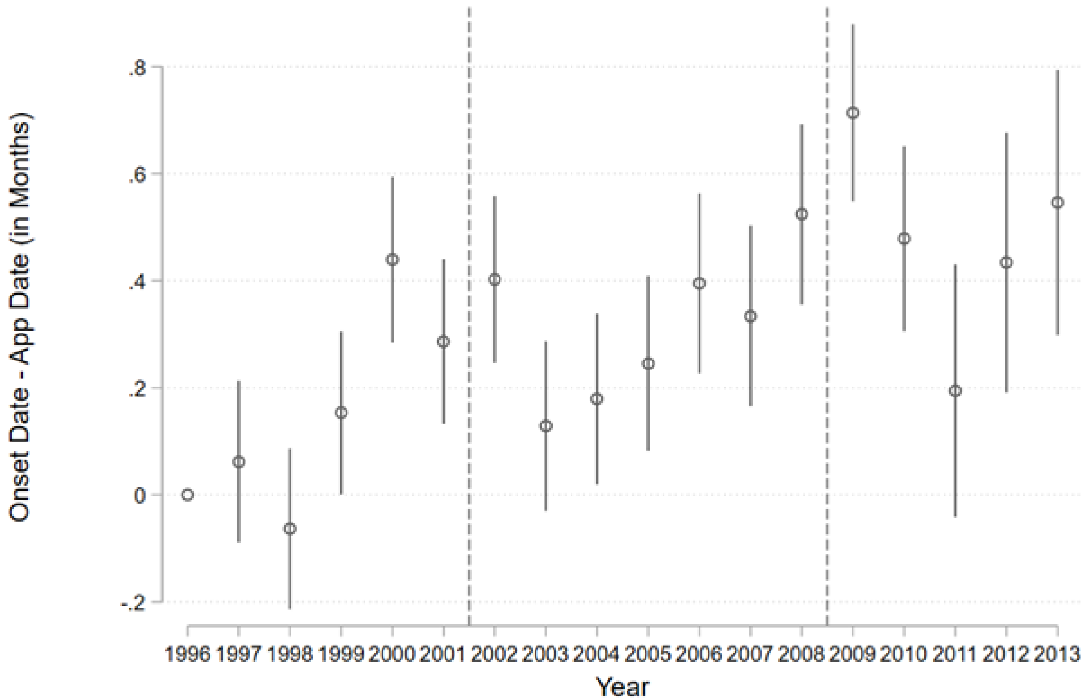

Figure A8: Successful Backdating: Onset Date Minus Application Date For High Benefit Claimants Relative to Low Benefit Claimants Since 1996

Notes: Sample restricted to claimants who applied between 1996 and 2013. Coefficients estimated from equation (4). Each coefficient represents the change in the number of days between the onset date and the application date for claimants with above $500 in monthly benefits relative to claimants with below $500 in monthly benefits. Sex, age at application, primary diagnostic code fixed effects, $100 monthly entitlement bin fixed effects, and year of application fixed effects are included. Since the onset date is typically before the application date, this is usually a negative number. An increase in the onset date minus the application date would be consistent with the onset date moving closer to the application date. We include a dashed line after 2001 because the 2002 change occurs in February 2002, and we include a dashed line after 2008 because the 2009 change occurs in June 2009. 95 percent confidence intervals based on robust standard errors are shown.

Source: Authors' own calculations from the SSA DAF-PUF.

Table A1. Summary Statistics, 1996-2002

	Less Affected	More Affected	More Affected, Excluding $1,500+	More Affected, Excluding $1,000+
Wait Time from Entitlement to Approval	13.2	9.22	9.27	9.85
	(12.9)	(11.1)	(11.1)	(11.4)
Hypothetical Attorney Fee	1086.3	1648.9	1640.1	1583.9
	(1124.0)	(1791.2)	(1769.7)	(1659.8)
Hypothetical Backpay	3347.5	6044.1	5996.0	5507.1
	(3695.2)	(8147.0)	(8023.7)	(6872.7)
Fraction Hitting $4000 Cap but Not $5300 Cap	0.023	0.071	0.071	0.067
	(0.15)	(0.26)	(0.26)	(0.25)
Fee Maximizing Wait Increase Under $5300 Cap	0.16	0.24	0.25	0.27
	(1.21)	(1.06)	(1.07)	(1.16)
Onset Date Before Application Date	0.94	0.93	0.93	0.93
	(0.24)	(0.26)	(0.26)	(0.26)
Onset Date Equals Application Date	0.040	0.052	0.052	0.052
	(0.20)	(0.22)	(0.22)	(0.22)
Onset Date After Application Date	0.022	0.021	0.021	0.021
	(0.15)	(0.14)	(0.14)	(0.15)
Primary Diagnosis = Autism	0.0036	0.00076	0.00077	0.00094
	(0.060)	(0.028)	(0.028)	(0.031)
Primary Diagnosis = Int. Dis.	0.080	0.016	0.016	0.020
	(0.27)	(0.12)	(0.13)	(0.14)
Primary Diagnosis = Mood Dis.	0.32	0.23	0.23	0.25
	(0.47)	(0.42)	(0.42)	(0.44)
Primary Diagnosis = Int. Systems	0.25	0.34	0.34	0.32
	(0.43)	(0.47)	(0.47)	(0.47)
Primary Diagnosis = Musc. Skeletal	0.18	0.18	0.18	0.18
	(0.39)	(0.38)	(0.38)	(0.39)
Primary Diagnosis = Infection	0.14	0.21	0.21	0.19
	(0.34)	(0.41)	(0.41)	(0.39)
Primary Diagnosis = Congenital	0.010	0.0090	0.0089	0.0089
	(0.10)	(0.094)	(0.094)	(0.094)
Primary Diagnosis = Unknown	0.016	0.017	0.017	0.017
	(0.13)	(0.13)	(0.13)	(0.13)
Female	0.59	0.40	0.41	0.46
	(0.49)	(0.49)	(0.49)	(0.50)
Age at Application	38.3	44.4	44.3	42.9
	(10.9)	(9.44)	(9.45)	(9.48)
Some College or More	0.12	0.22	0.22	0.19
	(0.32)	(0.41)	(0.41)	(0.39)
Regional Unemployment Rate	6.99	6.05	6.07	6.22
	(3.66)	(3.24)	(3.25)	(3.30)
Regional Employment to Population Ratio	55.8	62.2	62.1	60.6
	(14.2)	(12.8)	(12.8)	(12.9)
Observations	94985	284252	278907	209983

Notes: This table shows summary statistics for claimants who applied under the $4,000 cap. Column 1 shows the mean and standard deviation (in parentheses) for each variable and for claimants with AIME from $0-$500. Column 2 shows these same statistics for claimants with AIME above $500.

Table A2. Effect of 2002 Fee Ceiling Increase, No Demographic or Diagnosis Controls

	(1)	(2)	(3)
Panel A. Wait Time from Entitlement to Approval (in Months), 1996-2009			
More Affected x Post-2002	1.433***	1.432***	1.598***
	(0.0651)	(0.0652)	(0.0677)
Constant	8.235***	8.609***	9.451***
	(0.0490)	(0.0489)	(0.0540)
Observations	1,124,297	1,036,503	786,427
R-squared	0.045	0.042	0.034
Panel B. Hypothetical Representative Fee (in Dollars), 1996-2009			
More Affected x Post-2002	518.2***	513.7***	488.7***
	(6.466)	(6.492)	(6.843)
Constant	1,325***	1,307***	1,265***
	(6.032)	(5.988)	(6.237)
Observations	1,124,297	1,036,503	786,427
R-squared	0.080	0.082	0.092
Sample Restriction	NO	Excluding $1,500+	Excluding $1,000+

Notes: *** p<0.01, ** p<0.05, * p<0.10. Robust standard errors in parentheses. This table estimates the difference-in-differences model from equation (3) for the 2002 fee ceiling increase. We limit the sample to claimants who applied under the $4,000 cap or the $5,300 cap. We then define Post-2002 as equal to one if the claimants applied under the $5,300 cap and zero otherwise. In Panel A, the outcome variable is the claimant's wait time from entitlement to claim approval (in months). In Panel B, the outcome variable is the hypothetical representative's fee (in $) based on the claimant wait time, claimant benefit level, and the fee ceiling in effect at the time of application. For both panels, all specifications **exclude** demographic controls (sex and age) and primary diagnosis fixed effects. Column 2 excludes claimants with benefit amounts above $1,500 per month, and column 3 excludes claimants with benefit amounts above $1,000 per month.

Table A3. Effect of 2009 Fee Ceiling Increase on Claimant Wait Time and Hypothetical Representative Fee

	(1)	(2)	(3)	(4)
Panel A. Wait Time from Entitlement to Approval (in Months), 2002-2013				
More Affected	0.553***	0.334***	0.291***	0.199**
	(0.0796)	(0.0776)	(0.0785)	(0.0818)
Constant	18.16***	11.17***	10.70***	10.31***
	(0.0716)	(0.507)	(0.556)	(1.001)
Observations	1,145,400	1,145,400	991,716	685,114
R-squared	0.049	0.126	0.115	0.098
Panel B. Hypothetical Representative Fee (in Dollars), 2002-2013				
More Affected x Post-2009	179.7***	137.5***	115.2***	68.04***
	(8.273)	(8.237)	(8.367)	(8.894)
Constant	3,212***	1,663***	1,549***	1,520***
	(10.90)	(120.9)	(124.8)	(196.9)
Observations	1,145,400	1,145,400	991,716	685,114
R-squared	0.056	0.144	0.145	0.146
Primary Diagnosis FEs	NO	YES	YES	YES
Sample Restriction	NO	NO	Excluding $1,500+	Excluding $1,000+

Notes: *** p<0.01, ** p<0.05, * p<0.10. Robust standard errors in parentheses. This table estimates the difference-in-differences model from equation (3) for the 2009 fee ceiling increase. We limit the sample to claimants who applied under the $5,300 cap or the $6,000 cap. We then define Post-2009 as equal to one if the claimants applied under the $6,000 cap and zero otherwise. In Panel A, the outcome variable is the claimant's wait time from entitlement to claim approval (in months). In Panel B, the outcome variable is the hypothetical representative's fee (in $) based on the claimant wait time, claimant benefit level, and the fee ceiling in effect at the time of application. For both panels, column 1 includes controls for claimant sex and age and column 2 includes those controls plus primary diagnosis fixed effects. Column 3 includes the same controls as column 2 but excludes claimants with benefit amounts above $1,500 per month, and column 4 includes those same controls controls but excludes claimants with benefit amounts above $1,000 per month.

Table A4. Effect of 2002 Fee Ceiling Increase, Labor Market Controls

	(1)	(2)	(3)	(4)
Panel A. Wait Time from Entitlement to Approval (in Months), 1996-2009				
More Affected x Post-2002	0.852***	0.848***	0.857***	1.177***
	(0.0622)	(0.0632)	(0.0633)	(0.0656)
Constant	7.224***	9.811***	9.813***	10.05***
	(0.440)	(0.607)	(0.635)	(0.965)
Observations	1,124,297	1,099,143	1,012,850	767,929
R-squared	0.149	0.155	0.151	0.137
Panel B. Hypothetical Representative Fee (in Dollars), 1996-2009				
More Affected x Post-2002	426.4***	429.5***	429.0***	432.7***
	(6.257)	(6.356)	(6.366)	(6.656)
Constant	865.4***	1,181***	1,201***	1,347***
	(107.1)	(132.4)	(132.4)	(175.2)
Observations	1,124,297	1,099,143	1,012,850	767,929
R-squared	0.194	0.198	0.199	0.198
Primary Diagnosis FEs	YES	YES	YES	YES
Labor Market Controls	NO	YES	YES	YES
Sample Restriction	NO	NO	Excluding $1,500+	Excluding $1,000+

Notes: *** p<0.01, ** p<0.05, * p<0.10. Robust standard errors in parentheses. This table estimates the difference-in-differences model from equation (3) for the 2002 fee ceiling increase. We limit the sample to claimants who applied under the $4,000 cap or the $5,300 cap. We then define Post-2002 as equal to one if the claimants applied under the $5,300 cap and zero otherwise. In Panel A, the outcome variable is the claimant's wait time from entitlement to claim approval (in months). In Panel B, the outcome variable is the hypothetical representative's fee (in $) based on the claimant wait time, claimant benefit level, and the fee ceiling in effect at the time of application. For both panels: column 1 includes demographic controls (age and sex) and primary diagnosis fixed effects. Column 2 includes those controls as well as the annual gender by education by region employment to population ratio, region fixed effects, and education group fixed effects. Column 3 includes controls but excludes claimants with benefit amounts above $1,500 per month, and column 4 includes controls but excludes claimants with benefit amounts above $1,000 per month.

Table A5. Effect of Fee Increase in Regions with Below vs. Above Median Backlogs

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Wait Time from Entitlement to Approval (in Months), 1996-2009								
Post-2002	1.212***	2.972***	2.189***	2.326***	1.943***	2.534***	2.095***	2.420***
	(0.0816)	(0.0723)	(0.0814)	(0.0727)	(0.0803)	(0.0735)	(0.0783)	(0.0752)
More Affected x Post-2002	0.783***	0.764***	0.640***	0.834***	0.713***	0.780***	0.702***	0.792***
	(0.0911)	(0.0813)	(0.0911)	(0.0816)	(0.0899)	(0.0826)	(0.0876)	(0.0845)
Constant	4.661***	1.231	3.187*	2.400	3.672*	2.146	3.957**	1.508
	(1.809)	(1.872)	(1.817)	(1.892)	(1.899)	(1.809)	(1.764)	(1.955)
Observations	467,209	631,934	503,821	595,322	509,874	589,269	535,746	563,397
R-squared	0.147	0.158	0.146	0.157	0.147	0.157	0.146	0.158
Growth in Pending Cases from	2000-01	2000-01	2000-02	2000-02	2000-03	2000-03	2000-04	2000-04
Above/Below Median Growth	Below	Above	Below	Above	Below	Above	Below	Above

Notes: *** p<0.01, ** p<0.05, * p<0.10. Robust standard errors in parentheses. This table estimates the difference-in-differences model from equation (3) for the 2002 fee ceiling increase. We split the sample by region based on growth in pending SSDI cases from 2000-2001, 2000-2002, 2000-2003, and 2000-2004. Growth in pending SSDI cases is calculated at the region-level by aggregating state-level workload data released by SSA. Columns 1, 3, 5, and 7 estimate the model for regions with below median growth in pending cases and columns 2, 4, 6, and 8 estimate the model for regions with above median growth in pending cases. We limit the sample to claimants who applied under the $4,000 cap or the $5,300 cap. We then define Post-2002 as equal to one if the claimants applied under the $5,300 cap and zero otherwise. The outcome variable is the claimant's wait time from entitlement to claim approval (in months). For all specifications, we also include demographic controls (age and sex) interacted with region fixed effects, primary diagnosis fixed effects, and benefit bin fixed effects interacted with region fixed effects.

Table A6. Effect of 2009 Fee Ceiling Increase by Primary Diagnosis

Primary Diagnosis	Mood Disorder	Int. Systems	Musc. Skel.	Infection
	(1)	(2)	(3)	(4)
Panel A. Likelihood of Reaching Pre-Increase Cap but Not Post-Increase Cap, 2002-2009				
More Affected x Post-2009	0.0288***	0.0293***	0.0499***	0.0143***
	(0.000670)	(0.000692)	(0.000909)	(0.000682)
Constant	-0.000304	0.0223***	0.0650***	0.0124***
	(0.00144)	(0.00173)	(0.00312)	(0.00161)
Observations	189,222	219,473	188,613	115,198
R-squared	0.004	0.003	0.007	0.001
Panel B. Wait Time from Entitlement to Approval (in Months), 2002-2013				
More Affected x Post-2009	0.121	0.756***	0.182	0.255
	(0.152)	(0.160)	(0.149)	(0.225)
Constant	12.87***	16.12***	37.92***	10.73***
	(0.126)	(0.135)	(0.178)	(0.144)
Observations	269,792	338,962	321,975	168,692
R-squared	0.020	0.046	0.126	0.048

Notes: *** p<0.01, ** p<0.05, * p<0.10. Robust standard errors in parentheses. In Panel A, we limit the sample to claimants who applied under the $5,300 cap. We then define More Affected as equal to one if the claimant's monthly benefits exceed $500 and zero if not. The outcome variable is an indicator that a claimant reaches the $5,300 cap but would not have reached the $6,000 cap. Panel B estimates the difference-in-differences model from equation (3) for the 2009 fee ceiling increase. We limit the sample to claimants who applied under the $5,300 cap or the $6,000 cap. We then define Post-2009 as equal to one if the claimants applied under the $6,000 cap and zero otherwise. The outcome variable is the claimant's wait time from entitlement to claim approval (in months). All columns include demographic controls (sex and age). Columns 1-4 are limited to claimants with diagnoses of: mood disorder, int. systems, musculoskeletal, and infection, respectively.

Table A7. Alternative Plausible Definitions of Share Treated

Attorney's return to…	Share of Treated Cases
(2 Month Wait, $5300 Cap)>(2 Month Wait, $4000 Cap)	0.139
(4 Month Wait, $5300 Cap)>(4 Month Wait, $4000 Cap)	0.186
(6 Month Wait, $5300 Cap)>(6 Month Wait, $4000 Cap)	0.239
(8 Month Wait, $5300 Cap)>(8 Month Wait, $4000 Cap)	0.299
(10 Month Wait, $5300 Cap)>(10 Month Wait, $4000 Cap)	0.378
(12 Month Wait, $5300 Cap)>(12 Month Wait, $4000 Cap)	0.481

Notes: This table shows the fraction of treated cases in 2001-2002 in which an attorney would benefit more from a longer wait under the $5,300 cap than under the $4,000 cap. The first row shows the fraction of cases in which the attorney would benefit more from a 2 month wait under the $5,300 cap than under the $4,000 cap. The second row shows the fraction of cases in which the attorney would benefit more from a 4 month wait under the $5,300 cap than under the $4,000 cap.

Online Appendix B: Analysis of the Wait Time Distribution

We explore how claimant outcomes change in close proximity around the kink point in the potential representative pay. This method highlights any precise, strategic behavior that pushes wait times towards the fee threshold. Because some aspects of the application and appeal process are out of the representative's control, we might expect heaping rather than exact bunching. As see in Figure 2, marginal revenue for representatives will kink to zero at the point when the individual's past due benefits pushes representative fees over the threshold. Because the month of the kink varies with the monthly benefit amount, strategic bunching might not be detectable in the distribution of the raw number of months from entitlement to award (see Appendix Figure B1). This kink point will vary across individuals, by entitlement level.[38] For this reason, we calculate the difference between the realized number of full months between entitlement and allowance date and the number of months that would reach the threshold. In other words we calculate

$$Month\ Waited\ Past\ Kink_i = Months\ Entitlement\ to\ Allowance_i - \frac{Threshold_{yr} * 4}{Benefit\ Level_i} \tag{6}$$

The fee threshold will either be \$4,000, \$5,300, or \$6,000 depending on the year, which is multiplied by 4 to capture the 25 percent rule. In Figure B2 we see that this distribution is centered around -12 months, but there is still considerable mass at 0, where the wait time between entitlement and allowance pushes 25 percent of past due benefits past the fee ceiling. There is no obvious evidence of strategic bunching.

To identify strategic bunching, we need to observe a counterfactual distribution. Rather than approximating a counterfactual distribution using smooth polynomials as is sometimes done in the literature, we will exploit changes in the fee ceiling and use years prior to the threshold change as a counterfactual distribution. To determine how the kink in representative incentives affect claimant wait times, we will estimate how the density around the kink point associated with the 2002 threshold differs between 2002 and 2001, the year before the policy change. We construct the Months Waited Past Kink measure in 2001 and 2002 using \$5,300 as the applicable Fee Threshold in equation (6). As seen in Figure B3, the year-to-year distributions are similar, but there is excess mass in the months leading up to and following the kink point in the latter year relative to the proceeding year.

To formally estimate these difference we collapse the data into one-month "Months Waited Past Kink" bins by application year. In each bin we sum the number of claimants in that bin in each application year. We then calculate the percent of claimants in each application year in each one-month bin, to facilitate a

[38]As seen in the second panel of Figure B1, the distribution of wait times is compressed towards zero for higher entitlement claimants.

comparison across years, where the number of total claimants might fluctuate. We then estimate changes in the density within 60 months of reaching the kink in a regression format as follow

$$\begin{aligned} \textit{Percent of Claimants}_{mt} = \sum_{\tau=-60}^{60} \beta_\tau (\tau \textit{ months from kink}_m) * PostIncrease_t \\ + \delta PostIncrease_t + \phi_m + X_{mt}\Gamma + \varepsilon_{mt} \quad (7) \end{aligned}$$

The outcome of interest is the percent of claimants that applied in year t for whom the difference between realized wait time and the threshold was m months. Because the number of new claimants varies from year-to-year and we are comparing across years, we normalize by the total during the year rather than using the number of claimants. The main coefficients of interest are the β_τ, which trace out excess density in the post years, for each month-bin m relative to pre-years. We limit the years to those right around the threshold (2000, 2001, and 2002), in hope that the claimant cases will be the most similar, but include two pre-years to insure that estimated effects are not driven by anomalies in the counterfactual year.

We are interested in bunching right around the threshold where m equals zero, but include indicators for up to 60 months before and after the kink point to examine behavior throughout the distribution. All bins where the absolute value of m is less than 100 months are included (within 100 months of kink), so the β_τ coefficients are interpreted relative to the density changes in these regions of the distribution, over five years from the kink point. We include fixed effects for m, the number of month difference between realized wait time and the threshold, as well as average characteristics of claimants in the month/year bin, including sex, age, initial entitlement level, and primary diagnostic code shares. The fixed effects capture the average density in the bin across all years in the panel, meaning the β_τ coefficients capture the relative change in density for claimants that filed in the post year. Robust standard errors are reported. The identifying assumption is that the density of wait times would have remained the same in the application year after the policy change if the threshold change had not occurred. We also provide analogous results exploiting the 2009 fee ceiling increase, with the caveat that this occurred during the Great Recession.

Distributional Bunching Results

We plot the estimated wait time density change in 2002 relative to 2000-2001 with 95-percent confidence intervals in Figure B4. We see a significant increase in density in the 12 months before and after the Months Wait Past Kink threshold. In some month bins, we observe increases in the number of claimants of up to 0.2 percentage points. In some cases, this represents a 20 percent increase in the density. The increase in density remains significant, although much smaller in magnitude, up through 24 months. There is also a significant reduction in density between 42 and 12 months prior to the kink threshold. We observe much smaller changes further away from the kink, most of which are not significant. A segment of claimants in 2002 waited substantially longer than similar claimants in 2000-2001, before the fee threshold increase. Congruent with the difference in differences results, this empirical pattern is consistent with the increase in maximum representative compensation dragging out wait times.

When looking at the distribution around the 2009 threshold kink point there is also heaping, but it follows a slightly different pattern. As with the 2002 policy change, there is significantly more mass in the months just before the threshold, and significantly less mass in the proceeding months (42-18 months before the threshold). The magnitude of effects is about the same. However, unlike the 2002 policy change, we do not observe significantly more mass in the months following the kink point. The pattern turns precisely at the kink point and if anything there is less density past the kink. Although the empirical patterns differ slightly, they are both consistent with some claimants being pushed closer to the threshold and waiting additional months.

Because the bunching analysis reveals heaping rather than precise bunching, it is possible this is simply capturing a trending shift in the distribution over time. It would be concerning to see a similar shift in the distribution if we were to compare 2007 to 2005-2006 or any other combinations of years away from the policy change. We test this in an event study style framework.

Using the collapsed month-by-year data, we estimate how the percent of claimants in the -12 to 12 month window of the 2002 kink threshold evolves over time since 1996 as follows

$$\begin{aligned} \textit{Percent of Claimants}_{mt} = \sum_{\tau=1997}^{2008} \beta_\tau(-12 \leq m \leq 12) * (t = \tau) \\ + \delta(-12 \leq m \leq 12) + \alpha_t + X_{mt}\Gamma + \varepsilon_{mt} \quad (8) \end{aligned}$$

The outcome is the percent of total claimants in year t that fall in bin m where m is the number of months difference between entitlement to allowance wait time and the month the fee ceiling is reached. The main coefficients of interest are the β_τ which trace out across application years how the density changes for month bins within 12 months of the 2002 threshold relative to the other monthly bins. We control for the direct effect of being within 12 months as well as average characteristics of claimants in the cell and year fixed effects. The omitted reference year is 1996. Robust standard errors are reported.

We estimate a similar regression for the 2009 policy change, but look at changes in the density 0-18 months before the kink point as this is where the bunching is concentrated in 2009 relative to 2007-2008. This regression is as follows

$$Percent\ of\ Claimants_{mt} = \sum_{\tau=2003}^{2013} \beta_\tau(-18 \leq m \leq 0) * (t = \tau) + \delta(-18 \leq m \leq 0) + \alpha_t + X_{mt}\Gamma + \varepsilon_{mt} \quad (9)$$

In this specification data from application year 2002-2013 are included and 2002 is omitted as the reference year. In both cases we restrict the sample to exclude the other policy change. For many individuals, the policy changes lead to a difference that would cause overlap between the distance to the 2002 threshold and the 2009 threshold. This makes it impossible to interpret pre-trends prior to 2002 for the 2009 change and treatment effects after 2009 for the 2002 change.

These event study plots are provided in Appendix Figure B5. In both cases the pre-trends are flat, but begin to turn upward and become significant one to two years before the policy change. A slight pre-trend in the one to two year immediately proceeding the policy is not unexpected as we only observe application year, not agreement year. Since the policy change applies to all representative agreements approved after the policy date, and claimants can enter a representative agreement at any point in the application process, and agreements can include escalation clauses, it is possible, for example, that people who apply in 2000 or 2001 enter an agreement later and face the 2002 policy parameters.

For the 2002 policy change we see no changes in the density in the -12 to 12 month region of the distribution prior to 2000. In 2000 the point estimate is higher, but insignificant, but the effect is positive and significant in 2001. Then from 2002 on the effect is large and significant, consistent with approximately a 0.4 percentage point increase in the density per one month bin that gradually increases to about 0.9 percentage points.

For the 2009 policy change, we see no changes from 2002 to 2006, followed by an upward trend beginning in 2007, and continuing through 2009, the year of the change. This corresponds to a 0.5 percentage point increase, which increases to about 0.7 percentage points and then levels off.

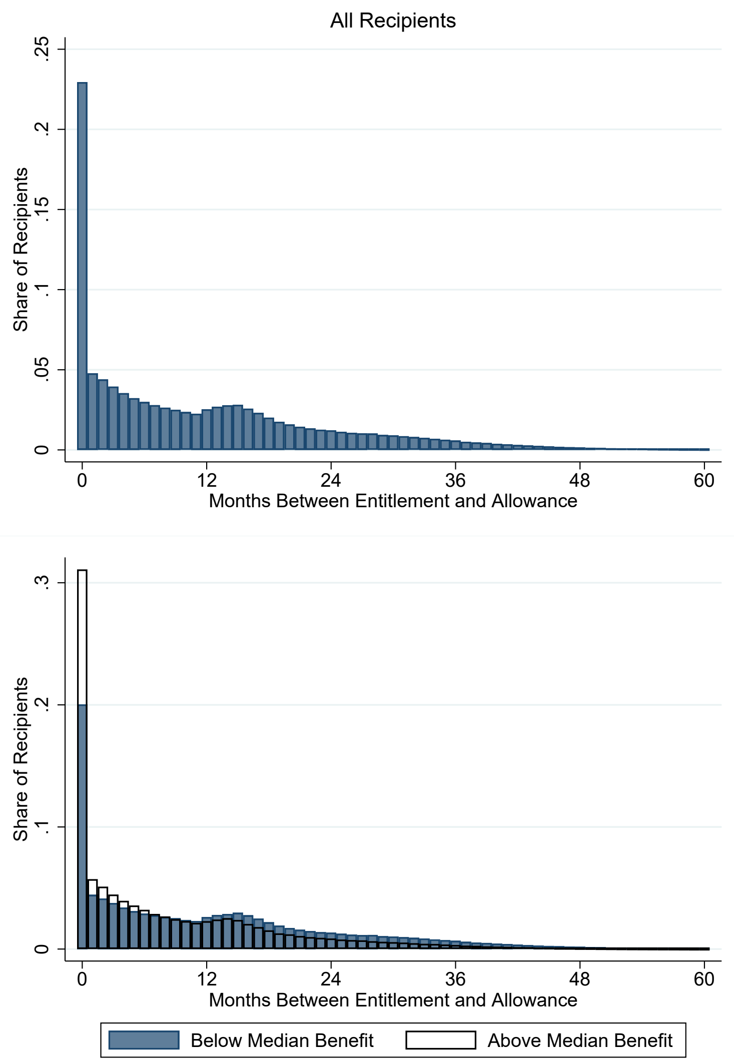

Figure B1: Distribution of Entitlement to Allowance "Wait Time" for Claimants

Notes: Claimant data from the DAF PUF. Sample restricted to primary claimants that only have one entitlement date, applied between 1996-2013, and had 60 months or less of wait time between entitlement to award dates. The share of claimants whose entitlement to allowance time (in full months) is plotted in one month bins.

Source: Authors' own calculations from the SSA DAF-PUF.

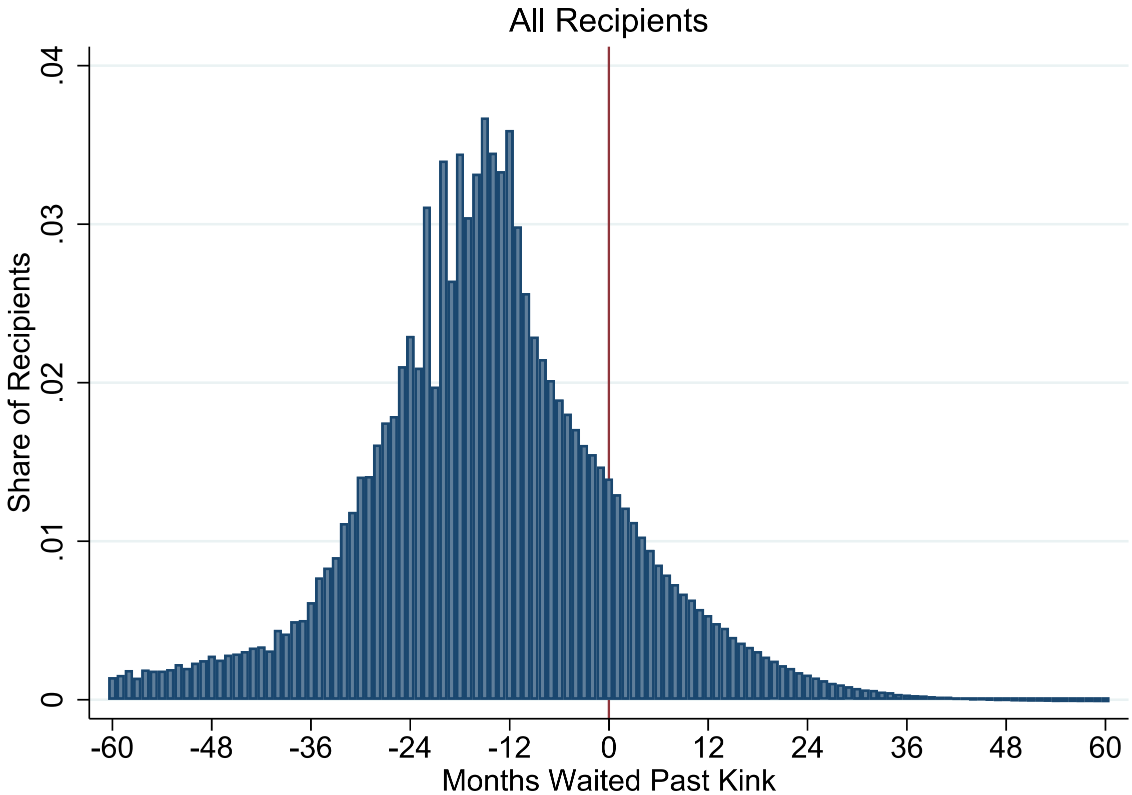

Figure B2: Distribution of Entitlement to Award "Wait Time" Relative to Prevailing Maximum Representative Fee Threshold

Notes: Claimant data from the DAF PUF. Sample restricted to primary claimants that only have one entitlement date, applied between 1996 and 2013, and the differences between wait times and time to the representative fee kink was 60 months or less. The share of claimants whose entitlement to award time (in full months) minus the number of months that would push past due benefits over the representative compensation threshold is plotted in one month bins. Threshold based on prescribed maximum representative compensation from the application year.

Source: Authors' own calculations from the SSA DAF-PUF.

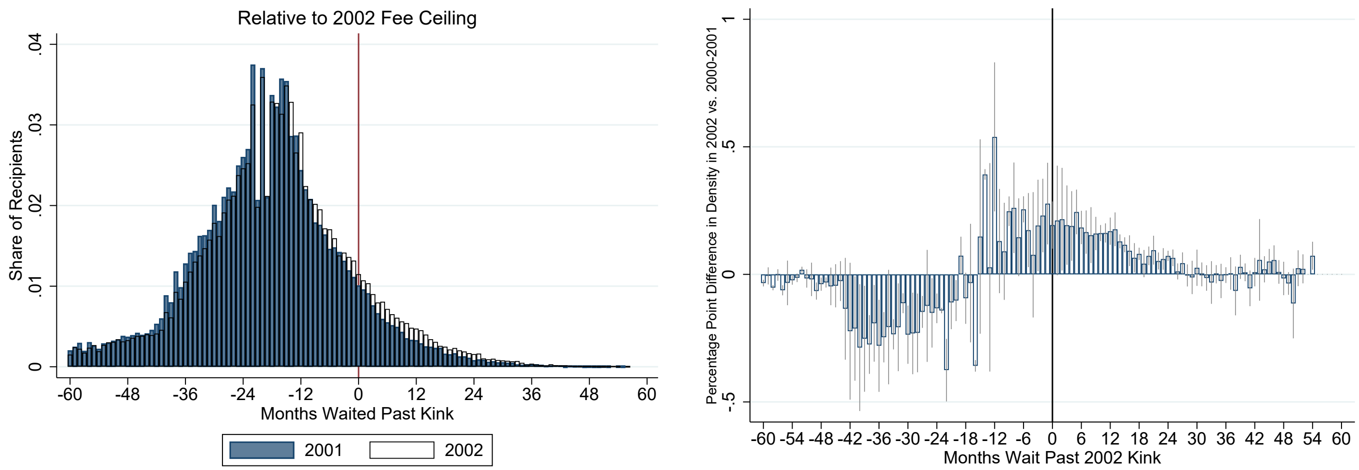

Figure B3: Distribution of Entitlement to Award "Wait Time" Relative to 2002 Maximum Representative Fee Threshold for Claimants Just Before and After Fee Ceiling Increase

Notes: Claimant data from the DAF PUF. Sample restricted to primary claimants that only have one entitlement date and the differences between wait times and time to the representative fee kink was 60 months or less. In the left panel claimants that applied in 2001 and 2002 are plotted, relative to the new, 2002 fee schedule. The share of claimants whose entitlement to award time (in full months) minus the number of months that would push past due benefits over the representative compensation threshold is plotted in one month bins. Threshold based on prescribed maximum representative compensation from the application year. In the right panel, only claimants that applied in 2000, 2001, or 2002 are included. Each bar is the interaction coefficient from equation (6) with 95 percent confidence intervals. We control for the month-by-year bin average gender, age, initial entitlement level, and primary diagnostic code shares.

Source: Authors' own calculations from the SSA DAF-PUF.

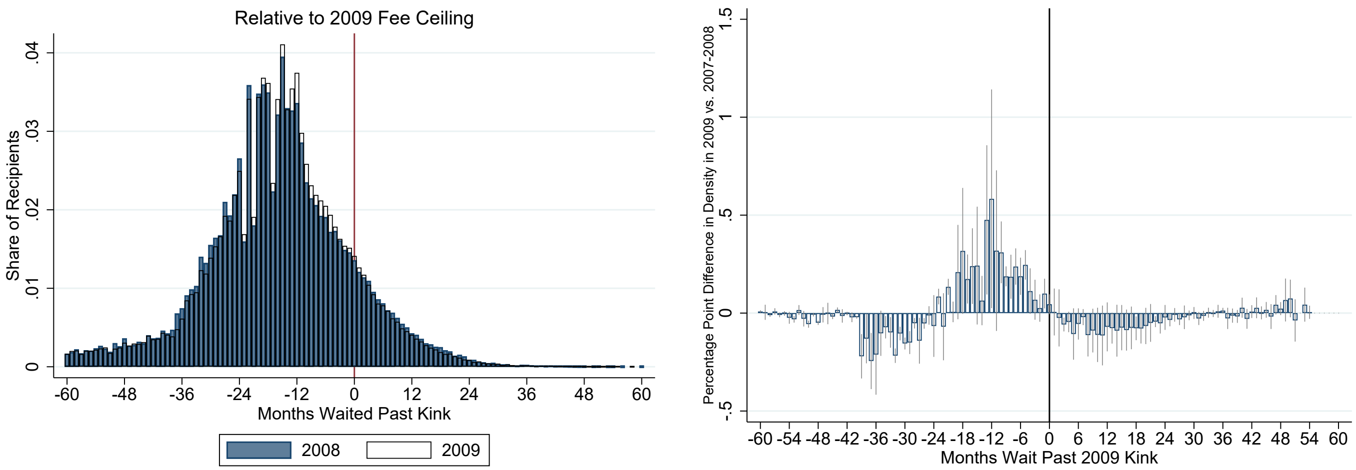

Figure B4: Difference in "Wait Time" Density Around New Fee Threshold in Pre- and Post Years

Notes: Claimant data from the DAF PUF. Sample restricted to primary claimants that only have one entitlement date. In the left panel only claimants that applied in 2008 and 2009 are plotted, relative to the new, 2009 fee schedule. The share of claimants whose entitlement to award time (in full months) minus the number of months that would push past due benefits over the representative compensation threshold is plotted in one month bins. Threshold based on prescribed maximum representative compensation from the application year. In the right panel only claimants that applied in 2000, 2001, or 2002 are included. Only claimants that applied in 2007, 2008, or 2009 are included in the right panel. Each bar is the interaction coefficient from equation (6) with 95 percent confidence intervals. We control for the month-by-year bin average gender, age, initial entitlement level, and primary diagnostic code shares.

Source: Author's own calculations from the SSA DAF-PUF.

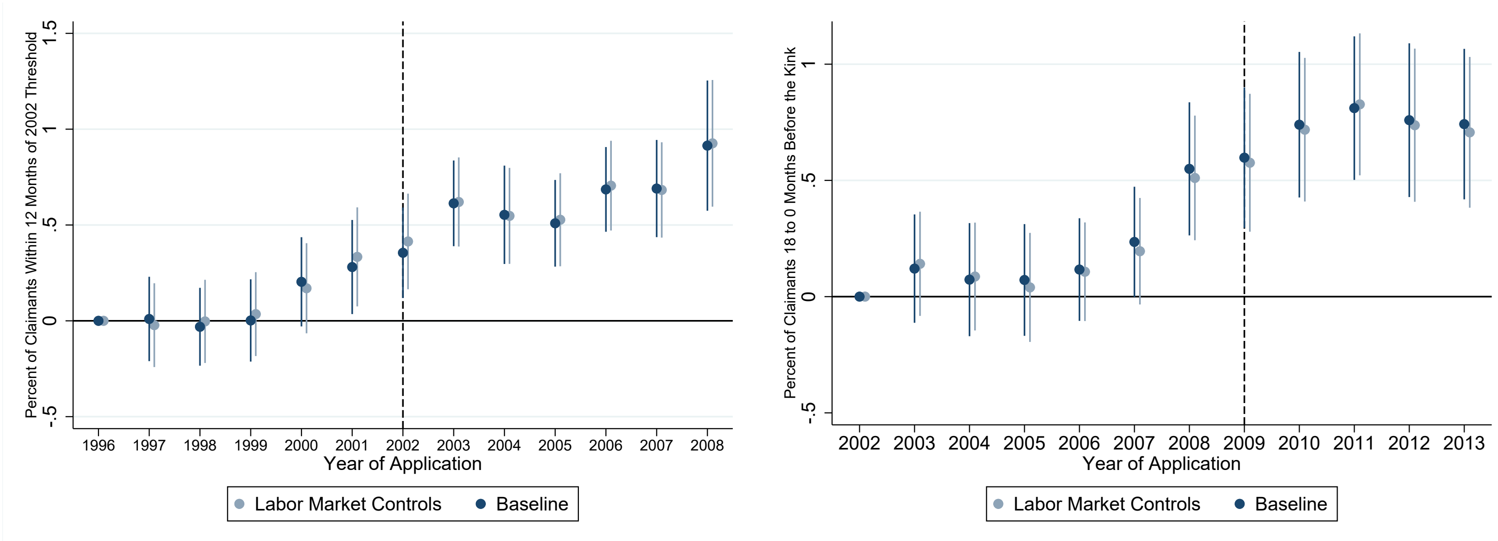

Figure B5: Event Study Difference in Bunching Around Fee Ceiling

Notes: Claimant data from the DAF PUF. Sample restricted to primary claimants that only have one entitlement date and applied in 1996 or later. Individual level data is collapsed to month-from-the threshold bins. The coefficients represent the change in the share of claimants in the specified bins relative to other parts of the distribution in each year with 95 percent confidence intervals. We control for the month-by-year bin average gender, age, initial entitlement level, and primary diagnostic code shares. The lighter plots also include education fixed effects, region fixed effects, and the average gender by education by region employment to population ratio in the bin and the share in each education and region bin to control for local labor market effects. Because the representative agreement date does not correspond to the application date, it is possible individuals who applied in the year leading up to the change face the higher fee ceiling.

Source: Authors' own calculations from the SSA DAF-PUF.